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For any remaining deficiencies the responsibility is of course my own.

Contents

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List of Abbreviations

ATE	Average treatment effect
2SLS	Two step least square
BuBA	Deutsche Bundesbank
CEE	Central and Eastern Europe
CPI	Consumer price index
DEM	Deutschmark
DEV	Developing countries
EUR	Euro
FDI	Foreign direct investment
FE-2SLS	Fixed effects-two step least square
GSOEP	German Socio-Economic Panel
IFS	International financial statistics
IND	Industrialised countries
IV	Instrumental variable
KS	Kolmogorov-Smirnov
L.P.	Levinsohn-Petrin
LR	Likelihood ratio
MiDi	Mikrodatenbank Direktinvestition
MNE	Multinational enterprise
OLS	Ordinary least square
O.P.	Olley-Pakes
R&D	Research and development
TFP	Total factor productivity
TRANS	Transition countries
Ustan	Unternehmensbilanzstatistik
WEU	Western Europe

Introduction

This thesis includes three self-contained essays in empirical economics. Two of them are concerned with the topic of Foreign Direct Investment (FDI) and its impact on domestic firms and their workforce. The third one contributes to the field of labour and health economics. All essays contain extensive empirical investigations based on micro data sets both at the firm-level and at the level of the individual.

In chapters 1 and 2, I take advantage of access to a unique data base that combines information on German firms' foreign operations (Mikrodatenbank Direktinvestition, MiDi; see Lipponer 2003) with their domestic balance sheets (Unternehmensbilanzstatistik, Ustan; see Deutsche Bundesbank 1998). Both sets of data are collected by the Deutsche Bundesbank (BuBa), Frankfurt. Within the context of a broader research project on the impact of outward FDI on domestic labour markets, Becker, Jäckle, and Muendler matched these data for the first time. The empirical estimations in chapter 3 are based on data included in the German Socio-Economic Panel Study (GSOEP; see SOEP Group 2001) at the German Institute for Economic Research, Berlin.

FDI and Domestic Firms. Only few people are enthusiastic about the expansion of national firms' operations to foreign locations. Indeed, most of the public and many politicians fear that affiliates abroad negatively affect the home economy. Managers, business leaders, and part of the press do not share this opinion. They consider the international expansion strategy of domestic firms an important channel towards enhancing competitiveness. I contributed to this debate by writing three papers. Two of them published in this dissertation. The third paper, which is joint work with Sascha O. Becker, Karolina Ekholm, and Marc Muendler (2005a), uses data on German and Swedish multinational enterprises (MNEs) to answer the following questions: What factors determine where multinationals choose to run their foreign affiliates? And how is the firms' employment in different locations affected by wages in those locations?

For multinationals from Sweden and Germany, the strongest predictors of location choice are host country GDP and geographical distance from the home country. A noteworthy difference between both sets of firms is that countries with highly skilled labour forces strongly attract German multinationals, while there is no evidence of such skill tracing for Swedish MNEs. Turning to the second question, we find that, given their respective location choices, German and Swedish firms exhibit similar responses of labour demands to international wage differentials. In MNEs from either country, affiliate employment tends to substitute for employment at the parent firm, where reactions are most pronounced with respect to wages in Western Europe (WEU). Moreover, we also find substitutability between parent and affiliate employment for subsidiaries in Central and Eastern Europe (CEE).[1] German MNEs that face a one percent higher wage at home are estimated to increase their employment in CEE by 2.2 percent. For Swedish MNEs, a one percent higher wage at the home market is associated with an employment increase at their affiliates in CEE by

[1]Since the wage gap between the home countries Sweden and Germany, on the one hand side, and WEU, on the other, are significantly smaller than the wage differentials between these countries and CEE the employment effects of the latter may be the economically more important.

1.8 percent.

While Becker, Ekholm, Jäckle and Muendler (2005a) test for the substitutability of labour across German parents and their foreign affiliates in different world regions, chapter 1 compares (purely) domestic firms and German MNEs, both prior to and after they have switched from national to multinational activities. At this juncture the following questions are raised: How much better are multinationals? Are successful firms more likely to invest abroad? And do newly founded MNEs grow faster than national companies?

Since the expansion decision in the firm is not brought about exogenously but depends on the situation at home as well as on business and production opportunities at foreign locations, estimates of the performance gap between switchers and national firms could be biased. This kind of problem necessitates applying an endogenous treatment approach. Therefore, using probit estimates of the decision to become an MNE, Heckman's (1978) treatment model is applied to account for potential endogeneity issues.

Chapter 1 provides empirical evidence on the following points: Good firms do become multinational enterprises. During the years before they invest in foreign countries, future MNEs already exhibit higher performance attributes in levels, i.e. they are larger in size, pay higher wages, produce with higher capital intensities, and they are more productive than future non-MNEs. These results are confirmed by several two-sample Kolmogorov-Smirnov (KS) tests. They show that distribution functions of all firm characteristics for nationals are dominated by those for switchers. With the exception of firm size, no significant ex-ante differences in terms of growth rates are found. Turning to a model of the decision to go multinational, it becomes clear that prior success, in terms of size, productivity, and a high portion of intangible assets in total assets, increase the probability of becoming an MNE.

Considering the time period after the regime change, I find that benefits of switching to the parent firm exist. Both productivity and wage growth are higher for newly founded MNEs than for national firms. Capital intensities at switchers evolve towards the use of capital, and the size of the home operations (measured by employment) is not affected.

In contrast to chapter 1, which looks at a wide range of MNEs' performance characteristics, the focus of chapter 2 lies on the question of whether the international diversification strategy of German firms is associated with a change in the skill intensity of their domestic workforce. Trade theory suggests that outward foreign direct investment to countries with an abundance of low-skilled workers compared to the domestic location leads to a decrease in the relative demand for unskilled labour. However, besides cost reductions, the expansion strategy of multinational enterprises is driven by market access motives. In fact, most of the work force of German MNEs' foreign affiliates is located in industrialised rather than developing or transition countries. In 2001, for example, German manufacturing MNEs employed 58 percent of their foreign workforce in high-income countries. In theoretical terms, operating subsidiaries in industrialised countries could either keep labour demand unaffected or shift it towards the more skilled.

Given these considerations, chapter 2 employs a sample of 1,557 German manufacturing MNEs between 1996 and 2001 to investigate the influence of offshore production on the domestic skill mix. I apply firm-level average wages at German-headquartered multinationals as approximation for the skill intensity at the parent operation and employ three different foreign activity measures: the ratios of (aggregated) foreign affiliate to domestic employment, output, and capital. Furthermore, I distinguish between the effects of FDI to transition, developing, and industrialised countries.

An important finding in chapter 2 is that foreign activities of German manufacturing MNEs are associated with higher average wages at domestic operations. I interpret this as evidence indicating that the international production strategy of German MNEs has affected the relative demand for high- and low-skilled workers at the parent firm. When distinguishing between different host regions, my results indicate that additional foreign affiliate employment in industrialised countries raises skill intensity. The same positive effect is found for subsidiaries in developing countries. In the case of transition countries, increases in foreign relative to domestic output carry higher average wages at the domestic location.

Health and Wages. The empirical investigations in chapters 1 and 2 correct for endogenous treatment variables and firm-specific, unobserved effects, respectively. Chapter 3 combines and extends the two methods by utilizing some recently developed estimation approaches proposed by Wooldridge (1995) and Semykina and Wooldridge (2005). In an application of these estimators to health and labour economics, chapter 3 examines the impact of health on wages for women and men in Germany.

There are a number of reasons why health and wages may be interrelated. First, improving health leads to an increase in productivity and hence in wages. Second, as Grossman (2001) states, if the returns to health investment increase with the salary, health should rise with wages, and the issue of reverse causality comes up. Apart from the latter, a number of further problems occur: First, as a self-reported health variable is used for estimation measurement error could induce biased coefficients. Second, panel attrition introduced by the endogenous decision to participate in the labour market – with one reason for it being the health status – may result in inconsistent estimation. And third, since the unobservable genetic endowment of each person is highly likely to be correlated with the individual state of health, the well known omitted variables bias problem arises.

In chapter 3, data from the GSOEP between 1994 and 2005 is used to estimate reduced form wage equations for women and men augmented by a variable measuring health satisfaction. Considering different estimation methods, I control for all of the above problems in one common framework. A number of tests provide evidence that for men selection corrections are indicated, while this issue does not cause any problems in the female sample. The results show that good health positively influences wages for both women and men. The health variable is found to be downward biased due to measurement error. For males, employing pooled OLS or 2SLS instead of methods accounting for selection and individual heterogeneity introduces an upward bias in the health coefficient.

References

Becker, Sascha O., Karolina Ekholm, Robert Jäckle, and Marc-Andreas Muendler, "Location Choice and Employment Decisions: A Comparison of German and Swedish Multinationals," *Review of World Economics*, 2005, *141* (4), 693–731.

Deutsche Bundesbank, "The Methodological Basis of the Deutsche Bundesbank's Corporate Balance Sheet Statistics," Monthly Report, 1998(10), 49-64.

Grossman, Michael, "The Human Capital Model," in Anthony J. Culyer and Joseph P. Newhouse, eds., *Handbook of Health Economics*, Vol. 1A, Amsterdam: Elsevier Science B.V., 2001, pp. 347–409.

Heckman, James J., "Dummy Endogenous Variables in a Simultaneous Equation System," *Econometrica*, 1978, *46* (4), 931–60.

Lipponer, Alexander, "A "New" Micro Database for German FDI," in Heinz Herrmann and Robert Lipsey, eds., *Foreign Direct Investment in the Real and Financial Sector of Industrial Countries*, Berlin: Springer, 2003, pp. 215–44.

Semykina, Anastasia and Jeffrey M. Wooldridge, "Estimating Panel Data Models in the Presence of Endogeneity and Selection: Theory and Application," *mimeo*, October 2005.

SOEP Group, "The German Socio-Economic Panel (GSOEP) after more than 15 years - Overview," in Elke, Holst, Dean Lillard, and Thomas A. DiPrete, eds., *Proceedings of the 2000 Fourth International Conference of German Socio-Economic Panel Study Users (GSOEP2000)*, Vierteljahreshefte zur Wirtschaftsforschung, 2001, 70(1), pp. 7-14.

Wooldridge, Jeffrey M., "Selection correction for panel data models under conditional mean independence assumption," *Journal of Econometrics*, 1995, *68*, 115–132.

Chapter 1

Going Multinational: What are the effects on home market performance?

A number of recent studies find evidence for the existence of a persistent performance gap between multinational enterprises (MNE) and their domestic competitors. This chapter investigates to what extent MNEs have superior performance characteristics, both prior to and after they have switched from national to multinational activities. In the first case results are quite clear: Future multinationals outperform domestic firms. When comparing ex-post performance of firms an endogenous treatment model is applied to account for selectivity issues. The results suggest that after switching, both productivity and wage growth are higher for newly founded MNEs than for national firms. Employment growth is superior before switching but does not exhibit significantly higher ex-post growth rates. Moreover, capital intensities at multinationals evolve towards the use of capital.

1.1 Introduction

Foreign direct investment (FDI) of domestic firms has attracted the interest of both the general public and politicians. The abrupt increase of multinational activities towards the end of the 20th century has raised concerns that domestic firms' foreign operations negatively affect home economies. Most managers and business leaders do not share this opinion. They consider the international expansion of domestic firms an important channel to enhance competitiveness. Economists can contribute to the heated political debate by evaluating performance characteristics of multinational enterprises (MNEs) relative to purely domestic firms. Since multinationals do not arise randomly selectivity issues need to be taken into account.

In this chapter, I investigate to what extent MNEs have superior performance attributes, both prior to and after they have switched from national to multinational activities. For this purpose the following questions are posed: How much better are multinationals? Are successful firms more likely to invest abroad? And do MNEs grow faster than national companies? To answer the first question I discuss differences between domestic firms, newly founded multinationals, and existing MNEs. The second and the third topics are covered by a comparison of new multinationals and national firms before, at the time of, and after switching. To assess a broad range of firm attributes, I have constructed five different performance measures: 1) firm size; 2) total factor productivity (TFP); 3) labour productivity; 4) average wage per firm; 5) capital intensities. Selectivity problems necessitate to apply an endogenous treatment approach for the evaluation of ex-post performance characteristics. Therefore, using probit estimates of the decision to become an MNE, Heckman's (1978) treatment model is used to account for potential endogeneity issues.

Theoretical predictions distinguish between ex-ante (before switching) and ex-post performance differences. If a firm decides to expand into foreign markets by employing FDI it needs to overcome legal, cultural, and social barriers. Only proficient firms are able to cope with these kinds of fixed costs and, thus, might self-select into foreign markets. Additionally, for large firms rewards to future expansions on the home market are low compared to smaller firms, whose next profitable project may be to increase their domestic market share. When turning to the home market performance after switching, theoretical models are less obvious. If there is no other alternative to serve foreign markets besides the set-up or acquisition of an affiliate (horizontal perspective), becoming an MNE would have no negative or even positive effects on domestic operations. If, in contrast, the purpose of a multinational's foundation is to vertically divide its production process, performance measures could rise or decline. Firm size, for example, is expected to decline if workers are laid-off due to cost-saving motives. An overall gain in competitiveness through cost reductions, on the other hand, may increase the number of employees at the parent location. The possible co-existence of market seeking and cost-reducing forces also makes predictions about domestic productivity growth ambiguous. Learning effects due to new technological and managerial inputs may play a positive role. Contrariwise, the efforts of restructuring a newly founded multi-plant enterprise could be accompanied by productivity losses at the domestic operation. Similar pros and cons can be discussed for all performance measures, and I proceed with a more extensive discussion of these issues in section 1.7.1. The crucial point, however, is that whether investing abroad improves home market performance or not is, in the end, an empirical question.

My findings provide evidence of the following points: During the years prior to the regime change, switchers exhibit higher performance attributes in levels, i.e. they are larger in size, pay higher wages, produce with higher capital intensities, and they are more productive. These results are confirmed by several two-sample Kolmogorov-Smirnov (KS) tests. They show that distribution functions of all firm characteristics for nationals lie to the left of those for switchers. With the exception of employment, firms that become MNEs do not grow faster than future non-MNEs in the three years before switching. After the regime change, both productivity and wage growth are higher for newly founded MNEs than for national firms. Capital intensities at multinationals evolve towards the use of capital, and switching does not affect the size of the home operation.

The remainder of this chapter begins with a brief summary of the existing literature. An overview of the data and a short discussion of the different performance measures is provided in section 1.3. Section 1.4 compares existing MNEs and switchers. Then I offer a detailed discussion of ex-ante differences in levels and growth rates, also including a set of Kolmogorov-Smirnov tests. The determinants of the switching decision are derived within a probit framework in paragraph 1.6, and ex-post performance differences are discussed in section 1.7. Section 1.8 concludes the chapter.

1.2 Related Literature

In a theoretical model, Helpman, Melitz and Yeaple (2004) describe the relationship between firm productivity and the engagement in different stages of international trade. Highly productive firms become multinationals (MNEs), less productive companies serve foreign markets by exports, and the least productive firms stay on their domestic markets. Based on these predictions, Girma, Kneller and Pisu (2005) present an empirical investigation for the UK using the concept of statistical dominance. They confirm that productivity distributions are ordered according to the Helpman et al. (2005) paper. Girma et al. (2005) also try to test for “marginal firms”, i.e. they evaluate productivity differences between first-time exporters and nationals, on the one hand, and newly founded, foreign owned MNEs (non-UK MNEs) and domestic producers, on the other. Their investigations provide some weak evidence that the productivity distribution of newly founded (foreign) MNEs dominates that of domestic exporters but they find no evidence of superior productivity for marginal exporters. In a similar study, Arnold and Hussinger (2005b) test the Helpman et al. (2005) setting for a sample of German manufacturers. Comparing the productivity distributions of purely national companies, domestic exporters, and firms with outward investment, their results exhibit support for the predictions of the theoretical model.

Earlier research mainly focuses on partial tests of the relationship between the different types of firms. Starting with the comparison of exporters and nationals, Clerides, Lach and Tybout (1998) ask whether learning by exporting is of importance. Their empirical investigations show that efficient firms become exporters, but they find no backward link between previous exporting activities and the firms’ cost structures. Similar studies are Arnold and Hussinger (2005a) and Girma and Greenaway (2002), who apply propensity score matching techniques. Arnold and Hussinger (2005a) show that productive firms enter export markets, but being an exporter has no significant effect on productivity gains at home. As an exception to other authors, Girma and Greenaway (2002) find for a sample of UK manufactures that exporting increases productivity. Bernard and Jensen (1999) analyse ex-ante and ex-post performance evolutions of newcomers on export markets. They find clear evidence that successful firms become exporters. Beyond other papers, Bernard and Jensen (1999) as well as the analysis at hand do not solely focus on productivity measures but use a wide range of performance characteristics. Moreover, in this study I consider MNEs instead of exporters and take the endogeneity of the investment decision into account.

Apart from the relation between exporting and serving domestic markets, other studies compare multinationals and domestic producers as well as multinationals and exporting firms. An example for the first case is Castellani and Navaretti (2004). Employing propensity score techniques for Italian manufacturers, the authors analyse the effect of FDI on parent firm characteristics. Their results suggest that foreign expansions improve the growth of productivity and output but exhibit no significant impact on employment. Egger and Pfaffermayr (2003) try to evaluate the investment behaviour of MNEs if they were purely exporting firms. They are searching for the counterfactual domestic investment to foreign activities. Using three different methods to account for the endogeneity of the FDI decision, Egger and Pfaffermayr (2003) show for a sample of Austrian manufacturers that foreign activities do not diminish domestic investment in intangible assets, while they increase investment in tangible assets.

1.3 Data and Construction of Performance Measures

In the study at hand, I use data from the USTAN (*Unternehmensbilanzstatistik*; see Deutsche Bundesbank 1998) data base at Deutsche Bundesbank (BuBa) between 1992 and 2001. Every firm in Germany that draws a bill of exchange in a given year is required by law to report its balance sheet to BuBa, which collects this information in its USTAN data base when the bill of exchange is rediscounted. The draft of bills of exchange remains a common form of payment in Germany. However, increases in BuBa's value threshold for reporting resulted in several drops of the sample and a decrease of the sample size over time. Table A.2 in the appendix of this chapter exemplifies the impact of the described sample reduction on the distribution of the variables employment and capital stock.[1] The table implies the existence of an attrition bias with respect to small companies, i.e. in the course of time mainly small firms drop out of the estimation sample. Among the variables extracted from USTAN are employment, firm age, investment, tangible and intangible assets, profits, intermediate input goods, sectoral definitions, wage bills, and regional information. All financial figures are deflated to unity at year end 1998 using the German CPI (from the IMF's International Financial Statistics).[2]

Information on German multinationals and their foreign affiliates is obtained from the MIDI database (Mikrodatenbank Direktinvestition; see Lipponer 2003) of the Deutsche Bundesbank. A firm is defined as a newly founded MNE (*switcher*) if the parent identifier appears in the MIDI dataset for the first time.[3] That is a multinational emerges if it "[...] acquires a substantial controlling interest in a foreign firm or sets up a subsidiary in a foreign country" (Markusen 2002, p.5). Firms from the MIDI database were string-matched by name to companies in the BuBa USTAN data set. The string matching routine automatically chose firms with an equality of at least 50 percent of all letters

[1]The table depicts summary statistics for the overall USTAN data set without any further adaptations.

[2]The end of 1998 is the mid point of the matched 1996-2001 FDI data (see below). In addition, the introduction of the euro in early 1999 makes December 1998 a natural reference date.

[3]A parent appears before 1999 if it controls at least 20 percent of its foreign affiliates' equity and the affiliates' balance sheet total is at least 1 million DM. After 1998 the affiliates had to satisfy either of the following two criteria: (i) the parent controls at least 10 percent of equity and the balance sheet total is at least 5 million EUR; or (ii) the parent controls at least 50 percent of equity and the balance sheet is at least 0.5 million EUR. Lipponer (2003) stresses that the modification of the notification limit in 1999 changes the number of reported affiliates significantly. However, as table 1.1 shows, the number of newly emerging parent firms (line 3) is not affected by this change.

included in their firm names. All potential matches were manually overseen before they were accepted as being the same company. Overall, a total of 2,955 unique firms were merged. I use information at the level of the foreign affiliate included in MiDi to construct (indirect) host country controls for the probit model of section 1.6. These information is aggregated over all existing MNEs per parent-sector and year. The currency conversion and deflation of foreign financial variables is described in appendix A.2.

Both the Ustan as well as the MiDi data set are available in the form of an unbalanced panel. I follow firms in the Ustan data base throughout the years 1994 to 2001. Individual parents in the MiDi data set are identifiable during the period 1996 to 2001. This allows to determine switchers between 1997 and 2001 and the comparison of ex-ante (before switching) parent characteristics between 1994 and 2000.[4] Table 1.1 summarises the development of the different data sets in the course of time starting in 1996. The first line reports the total number of Ustan firms for each year. In the second row the overall number of matched MiDi firms is depicted. These companies have already been MNEs in 1996 or switched status anytime between 1996 and 2001. A comparison with line five, which includes the total number of FDI firms in the MiDi data set, allows an evaluation of the matching algorithm. The merge process yields a matching quote between 18 percent in 2001 and 25 percent in 1997.[5] Line three reports firms that became multinationals between 1997 and 2001. Overall, 1,005 switchers appear in the matched sample. Row four reports the

[4]To avoid confusions about sample periods, a short note of clarification is presented at this point. For the purpose of evaluating firms I follow them throughout the period 1994-2001. In order to gain observations when constructing total factor productivity (see appendix), 1st step estimates of TFP refer to the time span 1992-2001, and 2nd step estimates refer to the period 1993-2001.

[5]Since for unmatched multinationals no parent information is available, performance attributes of matched FDI firms cannot be compared to characteristics of the overall number of multinationals in the MiDi data set.

remaining national companies of the USTAN data set.

Table 1.1: NUMBER OF FIRMS IN DIFFERENT DATA SETS, ALL SECTORS

	1996	1997	1998	1999	2000	2001	Total
USTAN total	69,423	62,341	48,194	41,102	36,207	26,737	284,004
Matched FDI firms	1,730	1,788	1,720	1,700	1,694	1,445	10,077
Matched switchers	-	272	210	232	201	90	1005
Nationals	67,693	60,553	46,474	39,402	34,513	25,292	204,504
FDI firms total	8,006	7,274	7,498	7,304	7,788	8,106	37,970

Source: USTAN and MIDI, Deutsche Bundesbank 1996-2001, own calculations.

My investigations are conducted at the firm level.[6] In order to eliminate parent firms founded for the mere purpose of acquiring or building up foreign affiliates, all parent companies that belong to the Nace 4digit sectors 6523 (other financial intermediation) and 7415 (management activities of holding companies) as well as companies with firm age below 5 years or firm size below 8 employees are removed from the estimation sample. Additionally, to prevent outliers from affecting results, variable values larger than the 99% and smaller than the 1% quantile were examined and if necessary dropped from the estimation sample. The large size of the USTAN sample allows for the use of Nace 4digit sector codes. However, for some estimations I have classified the firms into seven industry branches. Details of the aggregation can be found in table A.3 in the appendix.

The data at hand do not include information about export activities of firms. A domestic enterprise can therefore merely serve national markets or additionally be active on export markets. In this respect, another caveat is the missing possibility to identify domestic firms which are owned by foreign multinationals. In a recent study, Criscuolo and Martin (2003) find evidence for what they call the "MNE effect": MNEs, of foreign and domestic origin,

[6] I define firms as legally independent operations that draw a bill of exchange in a certain year.

are more productive than domestic firms. Either of these points might affect the results in this chapter in the same way. Switching premia calculated on different occasions could be downward biased since the comparison group goes beyond the definition of purely national firms in the above manner.

Five different firm attributes are employed in order to describe differences in the performance of switchers and nationals: 1) firm size; 2) total factor productivity (TFP); 3) labour productivity; 4) average wage per firm; 5) capital labour ratios.

As usual, firm size is measured by the number of employees. Total factor productivity is unobservable and needs to be estimated. The strategy in this study is to restrict technology parameters to a Cobb-Douglas production function and view the residual from the relationship between output and input factors as TFP. As is well known since the paper of Marschak and Andrews (1944), the correlation between unobserved, firm-specific productivity shocks and the firm's input choice causes a simultaneity bias.[7] In the literature different ways to deal with this problem have been documented. Following Olley and Pakes (1996) and Levinsohn and Petrin (2003), I use both investment in tangible and intangible assets and, in another specification, intermediate input goods as proxies to address the simultaneity problem. Consequently, three different TFP variables are constructed: a) TFP O.P. 1, using a semiparametric estimation approach, including regional dummies, a time trend, and applying Olley's and Pake's investment proxy; b) TFP O.P. 2, as a) but using firm age as an additional control variable;[8] c) TFP L.P., using the Stata ado file *levpet*

[7] In addition, if companies with smaller capital stock are more likely to close down their operations in consequence of a negative productivity shock, a selectivity problem occurs. In the USTAN data set firms drop out of the sample if they either exit the market or do not draw a bill of exchange in a certain year. Since it is not possible to distinguish these reasons, the selectivity issue cannot be addressed with the data at hand.

[8] In order to increase the number of observations, I did not employ firm age as explanatory

(see Levinsohn, Petrin and Poi 2003), which applies intermediate input goods as investment proxy. The appendix (A.1) includes a more detailed comparison of the different estimation methods.

In order to evaluate performance measures with respect to the firms' workforce, labour productivity, as constructed by the ratio of value added over employment, and the average wage per firm, measured as wage bill divided by employment, are used in the analysis. Finally, to assess capital intensities among different firms, capital labour ratios are used as a performance attribute. I measure capital labour ratios as the value of fixed assets (including real estate, machinery, and equipment; excluding immaterial and financial assets) divided by total employment.

1.4 How much better are Multinationals?

In this section, I discuss differences between newly founded multinationals and national firms, in the year of switching, as well as differences between existing multinationals and national firms. Performance gaps between both groups are calculated as percentage values in the following regression:

$$logP_{i,t} = \beta_0 + \beta_1 MNE_{i,t}(+\mathbf{c}_{i,t}\boldsymbol{\gamma}) + \delta_1 state_i + \delta_2 sector_i + \delta_3 year_t + u_{i,t}, \quad (1.1)$$

where $P_{i,t}$ depicts the corresponding performance measure, $MNE_{i,t}$ is a dummy variable that indicates multinational activities, $state_i$, $sector_i$, and $year_t$ refer to region, industry and time dummies respectively, and the vector $\mathbf{c}_{i,t}$ stands for the additional control variables firm age and size.[9]

Table 1.2 provides estimation results of the above equation. Each cell

variable in a).

[9]The dimension of the domestic operation is approximated by the number of employees. It is not included if the dependent variable, $P_{i,t}$, is firm size.

includes the coefficient of the $MNE_{i,t}$ dummy for another (dependent) performance variable. Columns (3) and (4) report the premia of already being an MNE between 1996 and 2001, whereas columns (1) and (2) describe the premia of becoming an MNE for the period 1997 to 2001. In the later case, all existing MNEs as well as switchers before and after the time of switching were removed from the estimation sample. In order to extract the maximum number of treated firms, I decided to assess different samples sizes for any of the separate performance regressions. Columns (2) and (4) depict results after adding additional controls ($\mathbf{c}_{i,t}$).

The performance gap for all firm attributes is positive and significantly different from zero.[10] Largest differences are found with respect to firm size. The number of employees at existing multinationals is 130% to 140% higher than at national firms. At the time of switching, newly founded MNEs are about twice as large as nationals. All productivity measures exhibit a persistent efficiency gap. Differences in total factor productivity range from 22% to 66%. As in the case of firm size, being an MNE is accompanied by a higher performance differential than becoming a multinational. This could be seen as first evidence for the existence of a positive performance dynamic after switching, i.e. becoming an MNE could have a positive impact on the post-investment productivity of parent firms. Performance measures related to the firms' work force show positive differences for all specifications. The average wage rate (labour productivity) at existing MNEs is 15% to 18% (25% to 26%) higher than at national firms. In the year of the regime change, MNE mark-ups for average wages (labour productivity) are between 11% and 13% (22% and

[10] To improve comparability between sectors I also constructed performance characteristics in deviation of the corresponding sector means. The use of these relative measures as dependent variables in equation (1.1) did not alter performance premia in any important way. Another consistency check was to construct equal sample sizes for each performance regression related to a certain column of table 1.2. Again, results did not change in an important manner.

Table 1.2: PERFORMANCE GAP, SWITCHERS VS. NATIONALS AND MNES VS. NATIONALS, ALL SECTORS

	Switchers[a)]	Switchers, ctrl.[b)]	MNE	MNE, ctrl.
	(1)	(2)	(3)	(4)
Employment	1.026[c)]	1.046	1.413	1.310
	(.052)	(.055)	(.016)	(.016)
TFP O.P. 1	.331	.222	.408	.261
	(.020)	(.021)	(.006)	(.006)
TFP O.P. 2	.341	.221	.411	.260
	(.022)	(.021)	(.006)	(.006)
TFP L.P.	.509	.222	.657	.271
	(.025)	(.021)	(.008)	(.006)
Labour productivity	.215	.246	.247	.264
	(.020)	(.022)	(.006)	(.006)
Average wage	.127	.106	.176	.147
	(.013)	(.014)	(.004)	(.004)
Capital/Labour	.210	.182	.330	.142
	(.057)	(.061)	(.017)	(.018)
N employment[d)]	99,487	93,561	186,572	167,740
	(690)	(551)	(7,782)	(7,043)
N TFP O.P. 1	94,544	89,396	177,130	159,953
	(643)	(513)	(7,041)	(6,372)
N TFP O.P. 2	89,396	89,396	159,953	159,953
	(513)	(513)	(6,372)	(6,372)
N TFP L.P.	94,544	89,396	177,130	159,953
	(643)	(513)	(7,041)	(6,372)
N labour prod.	98,615	92,926	184,936	166,583
	(670)	(538)	(7,668)	(6,945)
N avrg. wage	98,820	93,074	185,304	166,855
	(673)	(541)	(7,668)	(6,958)
N capital/labour	97,659	91,841	183,158	164,724
	(682)	(545)	(7,711)	(6,983)

Source: USTAN and MIDI, Deutsche Bundesbank 1996-2001, own calculations.
[a)] Switchers are observed in the first year of being a MNE. All existing MNEs as well as switchers before and after the time of switching were removed from the estimation sample.
[b)] Coefficients in columns (2) and (4) are estimated using firm age and firm size as additional control variables. In row (1) only firm age is used as an additional control variable.
[c)] Each cell includes the coefficient of the $MNE_{i,t}$ dummy for another performance variable in a separate regression. Standard errors are in parenthesis. If a parameter fails to be significant at the 10% level, it is set in italics.
[d)] Each N refers to the number of observations in the different performance regressions. The number of treated observations ($MNE_{i,t} = 1$) are set in parenthesis. Existing MNEs are observed in the time period from 1996 to 2001. Switchers are observed between 1997 and 2001.

25%). These differences may indicate a skill bias towards high skilled workers in the labour force of MNEs, which fits with the argument that being an MNE is accompanied by a shift in the firms' labour demand from production to non-production workers. Finally, I also investigate performance distinctions with respect to capital labour ratios. Differences in the capital intensities are between 18% and 21% for switchers and 14% and 33% for multinationals already active on foreign markets for a couple of years. This is consistent with the hypothesis that German firms tend to keep their more capital-intensive activities in Germany as well as with the hypothesis that MNEs tend to be more capital-intensive firms than non-MNEs.

Table 1.2 should not be misunderstood in view of a causal link between multinational activities and performance attributes at the domestic market. Rather, the results reveal positive correlation patterns that confirm inherent performance differences for a series of firm characteristics. It is shown that multinational enterprises exhibit superior performance features. Differences are even larger if firms were already active on foreign markets for a couple of years. The following sections investigate the performance premia of switchers in more detail.

1.5 Performance before Switching

Many studies have shown that multinationals outperform firms that serve only domestic markets.[11] For example, Barba Navaretti and Venables (2004) argue that "MNEs are larger and sometimes more productive than national firms." In a recent paper Helpman et al. (2004) state that heterogenous firms need to surpass certain productivity thresholds before they enter foreign markets. When companies start up or acquire affiliates abroad, they have to overcome a

[11] For a list of examples and references see Caves (1996) and Markusen (2002).

number of barriers to entry. Caves (1996) claims in his book that "the business firm [...] has a clear-cut national base and identity, with its internal planing and decision making carried out in the context of that nation's legal and cultural framework." That is, when investing abroad, firms need to deal with fixed costs due to legal, cultural, and social differences. Hence, it seems obvious that only firms with successful operations on domestic markets can handel the additional efforts that accompany the transformation into a multinational enterprise.

In this section two questions concerning ex-ante performance differences are assessed empirically: 1) Is there a performance gap (in levels) between switchers and nationals firms before switching? 2) What about performance growth in the run up to become an MNE? To back up these investigations, I conduct a set of Kolmogorov-Smirnov tests in section 1.5.2.

1.5.1 Ex-ante differences in levels

In section 1.4 evidence was found that multinationals at the time of switching have superior performance characteristics compared to their national counterparts. Consequently, the next step is to ask whether these differences also exist in the years prior to the regime change. To do so, the following equation is estimated:

$$\begin{aligned} logP_{i,T-t} &= \beta_0 + \beta_1 Switch_{i,T} (+\mathbf{c}_{i,T-t}\boldsymbol{\gamma}) \\ &+ \delta_1 state_i + \delta_2 sector_i + \delta_3 year_{T-t} + u_{i,T-t}, \end{aligned} \tag{1.2}$$

where T is the date of switching (1997-2001) and t determines the time lag ($t = 1, 2, 3$).[12] Performance attributes are assessed over a period of up to three years before switching. The corresponding time dimensions of the dependent

[12] All other variables are defined according to the covariates in equation (1.1).

variables in the estimation samples are therefore 1996-2000, 1995-2001, and 1994-2001.

Each cell of table 1.3 includes the coefficient of the $Switch_{i,T}$ dummy for a separate estimation of equation (1.2). Performance gaps for all firm attributes are significantly positive and – when taking confidence intervals on the estimators into account – roughly constant over time.[13] Again, firm size exhibits the largest differences. These are between 103% (98% without additional controls) in $T-1$ and 96% (91%) in $T-3$. Performance premia for switchers with respect to TFP range from values between 25% and 26% (35% and 53%) in the year before switching to dimensions of 21%-22% (30%-48%) three years before the regime change. Moreover, firms that become multinationals pay on average 10%-12% (11%-14%) higher wages, have a 23%-25% (22%) larger valued added per worker, and capital intensities exceed those of national firms by 15%-19% (22%-25%). A comparison with table 1.2 shows that performance differences, found in the year of switching, already existed in roughly the same magnitude up to three years before the firms become MNEs.

1.5.2 Kolmogorov-Smirnov tests on the equality of distributions

To back up the investigations in section 1.5.1 several Kolmogorov-Smirnov (KS) two-sample tests on the equality of performance distributions are conducted.[14] The tests provide an opportunity to determine differences in the distributions of firm attributes for switchers and non-multinationals. That is, they compare not only the first moments of the distribution functions but test

[13]Performance measures in deviation of the corresponding sector means and equal sample sizes did not alter results in any important way.

[14]These tests are implemented using the software package Stata. The KS test has no underlying distributional assumptions. It is therefore a non-parametric test. Additionally, t-tests on mean-differences were accomplished. They confirm the findings in the KS setting.

Table 1.3: PERFORMANCE GAP, FUTURE MNES VS. NATIONALS t YEARS BEFORE SWITCHING, ALL SECTORS

	Lag1[a)]	Lag1 ctrl.[b)]	Lag2	Lag2 ctrl.	Lag3	Lag3 ctrl.
	(1)	(2)	(3)	(4)	(5)	(6)
Employment	.976[c)]	1.030	.928	.995	.912	.956
	(.060)	(.061)	(.062)	(.062)	(.066)	(.064)
TFP O.P. 1	.353	.253	.333	.222	.308	.207
	(.024)	(.025)	(.024)	(.024)	(.025)	(.025)
TFP O.P. 2	.356	.252	.325	.221	.303	.207
	(.024)	(.025)	(.025)	(.024)	(.025)	(.025)
TFP L.P.	.534	.261	.508	.232	.480	.215
	(.029)	(.025)	(.029)	(.024)	(.030)	(.025)
Labour productivity	.221	.245	.217	.228	.223	.237
	(.023)	(.024)	(.024)	(.024)	(.024)	(.024)
Average wage	.138	.121	.121	.102	.113	.106
	(.015)	(.017)	(.015)	(.016)	(.016)	(.016)
Capital/labour	.241	.157	.223	.153	.252	.187
	(.066)	(.071)	(.068)	(.068)	(.070)	(.069)
N[d)] employment	92,504	88,070	83,572	80,103	75,273	72,361
	(492)	(441)	(458)	(421)	(408)	(386)
N O.P. 1	88,040	74,039	79,509	76,497	71,541	69,030
	(458)	(360)	(422)	(389)	(380)	(360)
N O.P. 2	84,180	74,039	76,499	76,497	69,032	69,030
	(409)	(360)	(389)	(389)	(360)	(360)
N L.P.	88,040	74,039	79,509	76,497	71,541	69,030
	(458)	(360)	(422)	(389)	(380)	(360)
N labour prod.	91,809	76,917	83,024	79,635	74,824	71,979
	(483)	(379)	(448)	(414)	(401)	(380)
N average wage	91,973	77,002	83,117	79,694	74,876	72,006
	(485)	(381)	(451)	(416)	(403)	(383)
N capital/labour	90,762	75,864	81,933	78,511	73,704	70,837
	(485)	(380)	(451)	(414)	(403)	(381)

Source: USTAN and MIDI, Deutsche Bundesbank 1994-2001, own calculations.
[a)] Performance characteristics of switchers are observed in the three years before switching ($T-1$-$T-3$).
[b)] Coefficients in columns (2), (4) and (6) are estimated using firm age and firm size as additional control variables. In row (1) only firm age is used as an additional control variable.
[c)] Each cell includes the coefficient of the $Switch_{i,T}$ dummy for another performance variable in a separate regression. Standard errors are in parenthesis. If a parameter fails to be significant at the 10% level, it is set in italics.
[d)] Each N refers to the number of observations in the different performance regressions. The number of treated observations ($Switch_{i,T} = 1$) are set in parenthesis. Performance measures of switchers are evaluated between 1994 and 2000. The formation of new MNEs is observed between 1997 and 2001. All existing MNEs as well as switchers before and after the time of switching were removed from the estimation sample.

whether the distributions (densities) of P_{T-t} with respect to newly founded MNEs are to the right of those for national firms.

Earlier papers written by Girma et al. (2005) and Arnold and Hussinger (2005b) use the above setting to check whether productivity levels of multinationals exceed those of exporters, which in turn are questioned to be grater than the productivity levels of purely national firms. Their studies accomplish a set of Kolmogorov-Smirnov tests on a contemporaneous basis, i.e. they ask whether *existing* MNEs are significantly different to nationals and exporters. Girma et al. (2005) also apply KS tests on a subset of first-time exporters in the period before they change export status and on a sample of foreign owned domestic firms in the year before they were acquired by foreign multinationals. Unlike the above studies, the analysis at hand tests for differences in parent characteristics between future multinationals and nationals up to three years before switching.

Table 1.4 provides results of the Kolmogorov-Smirnov tests for three time lags. The first block (rows (1) - (6)) refers to performance differences in $T-1$, the second one to $T-2$, and the third block to differences in $T-3$. The KS test makes use of the maximum vertical difference (D_{T-t}) between the distribution functions of switchers ($F(P^s_{T-t})$) and nationals ($F(P^n_{T-t})$). Rows (1), (7) and (13) include the largest positive deviations, $D^+_{T-t} = \max(F(P^n_{T-t}) - F(P^s_{T-t}))$, in the cumulative fractions of both groups. The corresponding p-values are reported in the lines below.[15] Thus, the hypothesis that the distribution function of a certain firm attribute P_{T-t} for nationals lies to the left of the distribution function for switchers is tested by asking whether P_{T-t} for nationals contains smaller values than for newly founded multinationals. Accordingly, maximum

[15] All p-values presented in table 1.4 are based on the asymptotic distributions derived by Smirnov (1939).

Table 1.4: KOLMOGOROV SMIRNOV TESTS OF THE EQUALITY OF DISTRIBUTIONS, ALL SECTORS

	Employment	TFP O.P. 1	TFP O.P. 2	TFP L.P.	Labour prod.	Average wage	Capital/labour
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Lag1, D^{+}_{T-1} [a)]	.448	.202	.245	.359	.181	.256	.11
p-value	0	0	0	0	0	0	0
Lag1, D^{-}_{T-1}	-.004	-.0003	-.0003	-.0006	-.003	-.0001	-.004
p-value	.986	1	1	1	.99	1	.987
Lag2, combined	.448	.202	.245	.359	.181	.256	.11
p-value	0	0	0	0	0	0	.00002
N[b)]	92,868	88,040	84,180	88,388	92.171	92,336	91,125
Lag2, D^{+}_{T-2}	.422	.213	.244	.366	.23	.238	.085
p-value	0	0	0	0	0	0	.001
Lag2, D^{-}_{T-2}	-.004	-.0002	-.0001	-.0006	-.0009	-.002	-.0007
p-value	.988	1	1	1	.999	.997	1
Lag2, combined	.422	.213	.244	.366	.23	.238	.085
p-value	0	0	0	0	0	0	.003
N	83,866	79,509	76,499	79,791	83,314	83,408	82,225
Lag3, D^{+}_{T-3}	.417	.178	.197	.331	.231	.229	.092
p-value	0	0	0	0	0	0	.001
Lag3, D^{-}_{T-3}	-.003	-.0008	-.002	-.002	-.001	-.002	-.0008
p-value	.99	1	.997	.998	.998	.996	.999
Lag3, combined	.417	.178	.197	.331	.231	.229	.092
p-value	0	0	0	0	0	0	.002
N	75,528	71,541	69,032	71,787	75,078	75,130	73,958

Source: USTAN and MIDI, Deutsche Bundesbank 1994-2001, own calculations.
[a)] Rows (1) + (2), (7) + (8), and (13) + (14) test whether the performance measures P_{T-t} for nationals contain smaller values than for switchers. Lines (1), (7), and (13) are the corresponding distances between the distribution functions. Rows (3) + (4), (9) + (10), and (15) + (16) test whether P_{T-t} contains larger values for nationals than for switchers. Lines (3), (9), and (15) are the corresponding distances between the distribution functions.
[b)] The number of observations refer to the overall number of firms in each test. The number of switchers equals the number of treated observations in the corresponding cells of table 1.3. Overall numbers of observations in table 1.3 are slightly lower since control variables are necessary when estimating the corresponding equations.

deviations in lines (3), (9) and (15) are defined as the statistic $D^-_{T-t} = \max(F(P^s_{T-t}) - F(P^n_{T-t})) = min(F(P^n_{T-t}) - F(P^s_{T-t}))$. These rows, together with (4), (10), and (16), test the hypothesis that P_{T-t} for nationals exhibits larger values than for switchers. Finally, row (5) of each block includes the combined test statistic $D = max(|D^+_{T-t}|, |D^-_{T-t}|)$.

Results depicted in table 1.4 confirm the findings in section 1.5.1.[16] In each of the three years before switching, national firms exhibit significantly smaller performance measures than future MNEs. For all firm characteristics, distribution functions for nationals lie to the left of those for switchers. The hypothesis that P_{T-t} for domestic producers exhibits larger values than for multinationals could be rejected. Furthermore, p-values of the combined test statistics are not higher than 0.003 and therefore reject the null hypothesis of the equality of distributions signifcantly.

1.5.3 Ex-ante differences in growth rates

At this stage a further question emerges. If level differences in performance attributes show premia for switchers, it seems natural to analyse deviations of performance growth rates in the run up to becoming a multinational. For that purpose the following regressions are estimated:

$$\begin{aligned}[logP_{i,T} - logP_{i,T-t}]/t &= \beta_0 + \beta_1 Switch_{i,T} (+\mathbf{c}_{i,T-t}\boldsymbol{\gamma}) \\ &+ \delta_1 state_i + \delta_2 sector_i + \delta_3 year_{T-t} + u_{i,T-t}. \quad (1.3)\end{aligned}$$

Growth rates are measured as yearly averages assessed over the three preceding years up to switching. So, the corresponding time dimensions of the dependent

[16] Implementing KS tests with performance measures in deviation of the corresponding sector means did alter vertical differences D_{T-t} slightly but had no impact on the overall results.

variables in the estimation samples are 1996-2001, 1995-2000, and 1994-1999.

The coefficient β_1 of the $Switch_{i,T}$ dummy measures average differences in growth rates per year between switchers and multinationals. Table 1.5 depicts results for the related time lags and all performance attributes. With regard to labour productivity, average wages, and the capital intensities average growth differences are close to zero, and non of the estimated parameters is significant. For the different TFP measures, some of the coefficients are found to exhibit significant values, but including additional controls takes away the effect from the treatment dummies. Yet, a significant correlation pattern is found with respect to the firm size. Depending on the time horizon, average employment growth is between two and five percentage points higher at future multinationals than at domestic companies.

These findings may be interpreted as evidence that firms preparing for a forthcoming expansion to foreign markets have additional personnel requirements. However, with respect to the other performance measures it seems that future MNEs – given they exhibit clearly higher performance attributes in levels – have already exploited most of their domestic growth potential.

My findings up to this point show clear differences between future multinationals and national firms. In the years prior to the regime change, switchers exhibit higher performance attributes in levels, they are larger in size, pay higher wages, produce with higher capital intensities, and they are more productive. Furthermore, the firms' size, as measured by the number of employees, grows faster at future multinationals than at national firms.

Table 1.5: DIFFERENCES IN PERFORMANCE GROWTH, FUTURE MNEs VS. NATIONALS t YEARS BEFORE SWITCHING, ALL SECTORS

	lag1[a]	lag1 ctrl.[b]	lag2	lag2 ctrl.	lag3	lag3 ctrl.
	(1)	(2)	(3)	(4)	(5)	(6)
Employment	.051[c]	.031	.043	.032	.024	.020
	(.009)	(.009)	(.006)	(.006)	(.005)	(.005)
TFP O.P. 1	*.017*	*.0005*	*.011*	*.006*	.016	.009
	(.011)	(.011)	(.007)	(.007)	(.005)	(.005)
TFP O.P. 2	*.009*	*.0002*	.012	*.006*	.014	*.008*
	(.011)	(.011)	(.007)	(.007)	(.005)	(.005)
TFP L.P.	.019	*.005*	.015	*.010*	.019	.012
	(.010)	(.010)	(.006)	(.006)	(.005)	(.005)
Labour productivity	-*.003*	-*.006*	*.004*	*.0008*	*.005*	-*.0005*
	(.010)	(.010)	(.006)	(.006)	(.006)	(.006)
Average wage	-*.006*	-*.013*	*.003*	*.001*	*.004*	-*.002*
	(.008)	(.008)	(.005)	(.005)	(.004)	(.004)
Capital/labour	*.002*	*.012*	*.014*	*.016*	*.010*	*.007*
	(.020)	(.021)	(.014)	(.015)	(.012)	(.013)
N[d] employment	90,947	86,648	81,610	78,266	73,189	70,399
	(490)	(439)	(457)	(420)	(406)	(384)
N TFP O.P. 1	86,155	82,461	77,093	74,260	68,947	66,619
	(453)	(404)	(417)	(384)	(372)	(352)
N TFP O.P. 2	82,461	82,461	74,260	74,260	66,619	66,619
	(404)	(404)	(384)	(384)	(352)	(352)
N TFP L.P.	86,155	82,461	77,093	74,260	68,947	66,619
	(453)	(404)	(417)	(384)	(372)	(352)
N labour prod.	90,093	85,940	80,910	77,668	72,591	69,885
	(477)	(430)	(443)	(411)	(396)	(376)
N average wage	90,322	86,118	81,063	77,782	72,675	69,949
	(478)	(431)	(444)	(412)	(396)	(378)
N capital/labour	89,071	84,837	79,783	76,492	71,416	68,675
	(482)	(431)	(449)	(412)	(400)	(378)

Source: USTAN and MIDI, Deutsche Bundesbank 1994-2001, own calculations.
[a] Growth rates are measured as yearly averages assessed over the three preceding years up to switching ($T-1$-$T-3$).
[b] Coefficients in columns (2), (4), and (6) are estimated using firm age and firm size as additional control variables. In row (1) only firm age is used as an additional control variable.
[c] Each cell includes the coefficient of the $Switch_{i,T}$ dummy for another performance variable in a separate regression. Standard errors are in parenthesis. If a parameter fails to be significant at the 10% level, it is set in italics.
[d] Each N refers to the number of observations in the different performance regressions. The number of treated observations ($Switch_{i,T} = 1$) are set in parenthesis. Performance measures of switchers are evaluated between 1994 and 2001. The formation of new MNEs (switching) is observed between 1997 and 2001. All existing MNEs as well as switchers before and after the time of switching were removed from the estimation sample.

1.6 The Decision to become a Multinational

A common approach that allows to take the endogeneity of the switching decision into account is the Heckman (1978) estimator. Applying this method necessitates to determine the factors behind the firms' decision to go abroad within a probit framework. Binary choice models suffer from the shortcoming that (foreign) location specific variables (host country attributes) are not identified.[17] It is, however, possible to apply indirect measures that control for host country effects. To construct such variables average foreign affiliate characteristics of existing MNEs are used. These attributes allow to augment the probit specifications with information on host country specifics of existing multinationals active in the same home market sector as potential switchers. Further details of this kind of control variables are discussed below.

In a recent study, Muendler and Becker (2006) estimate reduced-form location choice functions in order to control for selectivity issues in a multinational's location-specific labour demand. In this section, I present an adapted version of their first-step, location-choice model to explain driving forces behind the decision to become a multinational. It is assumed that in period $T-1$ a firm's management decides whether to become an MNE or not. Investing at foreign locations in period T means producing a vector of final goods $\mathbf{X}_{i,T} = (\mathbf{x}^H_{i,T}, \mathbf{x}^F_{i,T})$ at home ($\mathbf{x}^H_{i,T}$) and abroad ($\mathbf{x}^F_{i,T}$), whereas staying means serving foreign markets by exports or producing solely for the national market ($\mathbf{X}_{i,T} = \mathbf{x}^H_{i,T}$). For its switching decision the firm i maximises expected profits

$$E_{i,T-1}(\Pi_{i,T}) = E_{i,T-1}(\mathbf{p}_T \mathbf{X}'_{i,T} - c_{i,T}(\mathbf{X}_{i,T}, \mathbf{w}_T)), \tag{1.4}$$

[17] Econometric models that allow to control for country specific attributes are, for example, the conditional or the nested logit model. For a more detailed evaluation of this problem, see Becker et al. (2005a).

where $\mathbf{p}_T$ are final goods prices on competitive world markets, and $c_{i,T}(\cdot)$ is a firm's cost function depending on output $\mathbf{X}_{i,T}$ and a vector of (home and foreign specific) input prices $\mathbf{w}_T$.[18] Given the above optimisation problem, a firm's "switching-rule" can be written as

$$\text{Switch iff}: \; E_{i,T-1}[\Pi_{i,T}(\mathbf{x}_{i,T}^{F*}, \mathbf{x}_{1,i,T}^{H*}) - \Pi_{i,T}(\mathbf{x}_{i,T}^{F} = 0, \mathbf{x}_{2,i,T}^{H*})] > F_{i,T}, \tag{1.5}$$

where $F_{i,T}$ are sunk costs the firm faces when investing abroad, $\mathbf{x}_{i,T}^{F*}$ is the (optimal) part of the output vector that is produced in its foreign location(s), $\mathbf{x}_{1,i,T}^{H*}$ is the fraction of $\mathbf{X}_{i,T}^{*}$ produced at home in case of an investment abroad, and $\mathbf{x}_{2,i,T}^{H*}$ is the optimal domestic output in case no foreign affiliate(s) is (are) founded. Using equation (1.4) in (1.5) and adding a stochastic error term $\eta_{i,T-1}$ with zero mean and variance σ_η^2 yields:

$$S_{i,T} = \begin{cases} 1 \text{ if} & E_{i,T-1}[\mathbf{p}_T \mathbf{X}_{i,T}^{\prime *}] - E_{i,T-1}[c_{i,T}(\mathbf{x}_{i,T}^{F*}, \mathbf{x}_{1,i,T}^{H*}, \mathbf{w}_{i,T}) \\ & - c_{i,T}(\mathbf{x}_{i,T}^{F} = 0, \mathbf{x}_{2,i,T}^{H*}, \mathbf{w}_{i,T})] - F_{i,T} + \eta_{i,T-1} > 0 \\ 0 & \text{otherwise}, \end{cases} \tag{1.6}$$

where $S_{i,T} = 1$ means a firm decides to become an MNE, $E_{i,T-1}[\mathbf{p}_T \mathbf{X}_{i,T}^{\prime *}]$ are expected revenues from producing the optimal amount of output, and the second term on the right hand side of equation (1.6), $E_{i,T-1}[c_{i,T}(\mathbf{x}_{i,T}^{F*}, \mathbf{x}_{1,i,T}^{H*}, \mathbf{w}_{i,T}) - c_{i,T}(\mathbf{x}_{i,T}^{F} = 0, \mathbf{x}_{2,i,T}^{H*}, \mathbf{w}_{i,T})]$, depicts the cost benefits of producing abroad. As-

[18] In a more general model, one could allow for a flexible decision horizon that goes back further than one year. Additionally, the decision to invest abroad could depend on discounted future profit streams. In mathematical terms this would be

$$E_{i,T-t}\Big(\frac{\sum_{g=T}^{\tau} \Pi_{i,g}}{(1+r)^{\tau-T}}\Big) = E_{i,T-t}\Big(\frac{\sum_{g=T}^{\tau} \mathbf{p}_s \mathbf{X}_{i,g}' - c_{i,g}(\mathbf{X}_{i,g}, \mathbf{w}_g)}{(1+r)^{\tau-T}}\Big),$$

where r is the discount rate and τ is the point in time when a firm might close down its foreign operations.

suming $\eta_{i,T-1}$ to be standard normally distributed gives rise to a probit model, where the probability to switch is estimated as

$$P(S_{i,T} = 1) = P(S^*_{i,T} > 0) = P(\eta_{i,T-1} > -\mathbf{z}_{i,T-1}\boldsymbol{\alpha} - \mathbf{y}_{s,T-1}\boldsymbol{\theta}). \qquad (1.7)$$

Here, $S^*_{i,T}$ is a latent variable (e.g. the propensity to start an MNE) and the vectors $\mathbf{z}_{i,T-1}$ and $\mathbf{y}_{s,T-1}$ stand for the firm's expectations in period $T-1$ with regard to the decision rule of equation (1.6); $\mathbf{z}_{i,T-1}$ exhibits variation on the firm level, whereas $\mathbf{y}_{s,T-1}$ varies only over sectors, $\boldsymbol{\alpha}$ and $\boldsymbol{\theta}$ are the corresponding parameter vectors. The time period under consideration is 1997-2000. All existing MNEs as well as switchers before and after the time of switching were removed from the estimation sample.

To approximate expected revenues when producing abroad, I use the log of average affiliate turnover domestic competitors realise in their foreign locations in the year before switching, i.e. the average revenue MNEs active in the same home market sector as potential switchers make abroad in period $T-1$. In order to account for the expected cost benefits, $E_{i,T-1}[c_{i,T}(\mathbf{x}^{F*}_{i,T}, \mathbf{x}^{H*}_{1,i,T}, \mathbf{w}_{i,T}) - c_{i,T}(\mathbf{x}^{F}_{i,T} = 0, \mathbf{x}^{H*}_{2,i,T}, \mathbf{w}_{i,T})]$, lagged parent firm characteristics like different productivity measures, firm size (*ln* employment), *ln* liabilities/total assets, *ln* capital/labour ratio, *ln* equity, and *ln* average wage are used. Additionally, as a sector-specific control variable, the log of average wages domestic competitors pay in foreign countries are included. Since sunk costs cannot be directly measured, they are approximated by employing the number of existing MNEs from the same sector in period $T-1$. To account for the firm's innovative abilities, the log of its intangible to total assets ratio is included. For the purpose of controlling intra-sector market power, the proportion of each firm's value added to sector-wide value added is used. Finally, in most specifications firm age serves as an additional control.[19] Apart from the value-added ratio, all sector-specific variables refer to NACE 2-digit codes. Time dummies control for the foundation of MNEs in different years. All explanatory variables are lagged one period.

Since results of the probit estimates are of main interest with respect to the Heckman (1978) estimator applied in section 1.7, only the most important findings are briefly discussed at this point. Estimation results are depicted in table 1.6. Each specification refers to another lagged productivity measure and is used as selection equation for one of the different dependent variables of equation (1.9). It becomes clear from either specification that size and productivity in $T-1$ are important determinants of the choice to become an MNE, i.e. an increase of these attributes raises the probability to switch. Hence, findings in section 1.5 – large and productive firms go multinational – are supported.

In line with the existing literature, I find that domestic firms with large intangible to total assets ratios are more likely to run business abroad than firms staying on national markets. Intangible assets are supposed to have public good characteristics within multi-plant companies. Markusen (2002) generally names these kind of assets "knowledge capital". The particular characteristics of the knowledge capital (transportability, jointness, skill intensity) should help companies to overcome potential sunk costs. At this point, my findings – a high rate of intangible assets increases the probability to switch – confirms Markusen's theory.

Surprisingly, I find a positive correlation between average wages domestic competitors pay at their foreign locations and the probability to go multinational. Two arguments may solve this puzzle: First, assuming that the main

[19]In some specifications also firm age squared was included. However, standard t-tests on the influence of this variable rejected any significant (non-linear) effect of the age variable.

Table 1.6: PROBIT ESTIMATES OF THE PROBABILITY TO BECOME A MNE, ALL SECTORS, 1997-2000

	(1)	(2)	(3)	(4)
lag log (l.l.) TFP O.P. 1	.	.113[a)] (.048)**	.	.
(l.l.) TFP O.P. 2	.178 (.055)***	.	.	.
(l.l.) TFP L.P.	.	.	.	.140 (.057)**
(l.l.) labour productivity	.	.	.351 (.063)***	.
(l.l.) employment	.125 (.035)***	.112 (.032)***	.177 (.028)***	.100 (.034)***
(l.l.) liabilities/tot. assets	-.251 (.083)***	-.304 (.073)***	-.282 (.081)***	-.243 (.083)***
(l.l.) equity	.037 (.022)*	.038 (.020)*	.026 (.021)	.040 (.022)*
(l.l.) capital/labour	-.013 (.023)	-.010 (.022)	-.055 (.023)**	-.018 (.023)
(l.l.) foreign wages	.281 (.141)**	.372 (.132)***	.300 (.132)**	.391 (.138)***
(l.l.) foreign turnover	-.050 (.025)**	-.033 (.022)	-.038 (.022)*	-.035 (.024)
lag MNE count sector	-.0001 (.00007)	-.0001 (.00006)*	-.00009 (.00006)	-.00008 (.00007)
lag firm age	-.002 (.0006)***	.	-.002 (.0006)***	-.002 (.0006)***
(l.l.) intang. ass. /total ass.	.048 (.015)***	.061 (.014)***	.040 (.015)***	.050 (.015)***
(l.l.) firm val. add./sec. val. add.	.042 (.024)*	.031 (.022)	.	.040 (.024)*
(l.l.) average wage	-.018 (.092)	-.003 (.082)	-.157 (.095)*	.007 (.092)
West/East Germany	.105 (.092)	.072 (.083)	.042 (.089)	.039 (.089)
constant	-4.895 (1.338)***	-6.000 (1.284)***	-7.331 (1.223)***	-6.717 (1.379)***
year dummies	yes	yes	yes	yes
N	41879	44401	42417	41879
pseudo R2	.075	.07	.072	.073

Source: USTAN and MIDI, Deutsche Bundesbank 1997-2000, own calculations.
[a)] Standard errors are in parenthesis: * significant at ten, ** at five, and *** at one percent.

motive behind investing abroad is the access to other countries' markets (horizontal motive), high wages in foreign locations could simply reflect the fact that most FDI goes to places which are similar to Germany in relative factor endowments. Secondly, since Blomström, Fors and Lipsey (1997) for Sweden and Marin (2004) for Germany and Austria report evidence that MNEs locate skill-intensive activities abroad my results may indicate a skill-seeking motive behind German foreign direct investment.[20]

Finally, another interesting point is the negative influence of credit capital (short and long run liabilities/total assets) on the probability to become a multinational. This could both be an indicator for the negative impact of credit constraints, on the one hand, and – through the different financing structure of small, medium, and large firms in Germany – simply be another criterion for the size of an operation.

1.7 Performance after Switching

The next issue at hand is to investigate firm performance in a post-investment framework. The question tackled at this juncture is what happens to the efficiency of firms in the three years after their choice to become an MNE.

1.7.1 Theoretical considerations of post-investment developments

Theoretical answers to this question are not clear cut. Concerning firm size, it depends on whether the parent retains operations at home that are complementary or substitutional to foreign activities. A substitutional relationship,

[20]To investigate this problem in more detail one needs access to both, the skill structure of foreign subsidiaries and the skill distribution at the German parent firms. Unfortunately such information is not available in the BuBa MiDi and Ustan data sets.

which is likely when cost-saving reasons play a decisive role (vertical FDI), comes along with smaller operations on the home market. Contrariwise, even for purely cost-reducing FDI an employment gain at the domestic operation is possible if rewards to potential cost reductions allow the firm to increase its overall market share. When investing in overseas or/and industrialised countries, aspirations for better market access and the proximity-concentration trade-off need to be considered (horizontal FDI).[21] In this case, instead of exporting goods MNEs produce at the foreign location. Thus, the employment effect is twofold: On the one hand, the home operation could be larger if the firm exported goods to the host country. On the other hand, if there were no other opportunity to serve the foreign market besides the set-up of a foreign affiliate, becoming an MNE would have no negative or even positive effects on domestic firm size. Since, in reality, the decision to become an MNE is possibly brought about by the co-existence of both cost-reducing and market-seeking motives, the overall effect on the parent firm is ambiguous.[22]

Another effect I am interested in is the impact of the switching decision on productivity. Again, different theoretical aspects should be considered in this respect. One argument for productivity increases at the domestic firm is the public good characteristic of firm-specific assets. Pfaffermayr (1999) tests for a sample of Austrian manufacturers whether the volume of foreign output influences labour productivity at home through multi-plant scale effects. He finds that production at subsidiaries, Austrian firms run abroad, increases the productivity at domestic plants. Barba Navaretti and Venables (2004) argue that also changes in the composition of factor inputs and learning ef-

[21]In fact, a major part of German MNEs' foreign operations is concentrated in high- rather than low-income countries.

[22]Becker et al. (2005a) test for substitutability of labour in different world regions and Germany. The study conducts analysis for existing MNEs (long- or medium-term perspective) and finds for both industrialised regions (e.g. Western Europe) and for transition countries (Central and Eastern Europe) a substitutional relationship.

fects (technological and managerial knowledge) play a role. Since technological and managerial knowledge exhibit public good characteristics within firms, it seems obvious that learning through switching – whatever motives (vertical, horizontal, or both) are behind the decision – should positively affect domestic productivity. In case of changes in the composition of factor inputs, it is a priori not clear whether home market productivity gains or loses from the decision to found an MNE. Such changes are likely to occur if the management vertically divides the production process, meaning that labour intensive production stages are shifted abroad. However, whether in that case efficiency at the remaining operation increases or decreases cannot be predicted. Marin (2004) argues that Austrian and German firms take advantage of cheap and abundant high skilled labour in Eastern and Central Europe. Hence, in this situation the productivity evolution at the domestic location might suffer. Moreover, the tremendous efforts of restructuring a newly founded multi-plant enterprise may – at least in the short run – be accompanied by productivity losses at the domestic location.

Finally, I assess the development of average wages. On the firm level, increasing productivity should raise wages. The export of labour intensive, blue collar jobs to low-income countries also implies the rise of average wages. Moreover, horizontal FDI, i.e. replicating domestic operations at different locations in order to gain access to new markets, could either keep labour demand unaffected or shift it towards the more skilled and hence increase average wages.

To draw a conclusion on the above considerations, the effect of switching on firms' size, productivity, average wages, and capital intensities is theoretically ambiguous and it is therefore inherently an empirical question to explore how performance differences evolve after the rise of a new multinational.

1.7.2 Empirical considerations of post-investment developments

The easiest way to evaluate the effect of a regime change on parent firm attributes is to run simple OLS regressions of the firms' average outcome changes in $T + t$ on the switching status and a number of initial control variables in period T:

$$\begin{aligned} [logP_{i,T+t} - logP_{i,T}]/t &= \beta_0 + \beta_1 Switch_{i,T} + \mathbf{c}_{i,T}\boldsymbol{\gamma} \\ &+ \delta_1 state_i + \delta_2 sector_i + \delta_3 year_T + u_{i,T}, \quad (1.8) \end{aligned}$$

where T is the date of switching (1997-2000) and t is the time span we look ahead (t=1,2,3). Average performance growth is assessed over a period of up to three years after the decision to go multinational. The corresponding switching-dates for the different time spans are therefore 1997-2000, 1997-1999, and 1997-1998. Additional to firm size and firm age the vector $\mathbf{c}_{i,T}$ includes the average wage per firm over the average sector wage (not included if the dependent variable is average wage growth) and the value added per firm over the sector wide value added (not included for value added over employment). These variables are meant to control for the initial skill level of the firms' work force and the competitive position within the domestic sector.

As argued in the previous sections, it is likely that endogeneity issues bias results. Caves (1996) stresses that a typical firm, which decides to become a multinational, has already exploited most of its domestic growth potential. To the contrary, companies, that stay on domestic markets, may still have the potential to increase their home market performance without investing in foreign countries. Hence, one might suspect that firms, which are more likely to switch, exhibit on average lower performance growth rates (before switching) than their national counterparts, and thus the coefficient β_1 in equation (1.8) may be downward biased. The empirical descriptions in the previous sections verify this conjecture. While future MNEs clearly outperform national firms in levels, there exist – with the exception of employment – no performance differences in growth rates. Hence, equation (1.8) is re-estimated using the Heckman (1978) approach:

$$\begin{aligned}
[lnP_{i,T+t} - lnP_{i,T}]/t &= \beta_0 + \beta_1 Switch_{i,T} + \mathbf{c}_{i,T}\boldsymbol{\gamma} + \delta_1 state_i + \delta_2 sector_i \\
&+ \delta_3 year_T + \rho\sigma_\epsilon[\frac{\phi(\mathbf{z}_{i,T-1}\boldsymbol{\alpha} + \mathbf{y}_{s,T-1}\boldsymbol{\theta})}{\Phi(\mathbf{z}_{i,T-1}\boldsymbol{\alpha} + \mathbf{y}_{s,T-1}\boldsymbol{\theta})}] + \epsilon_{i,T} \quad (1.9) \\
S^*_{i,T} &= \mathbf{z}_{i,T-1}\boldsymbol{\alpha} + \mathbf{y}_{s,T-1}\boldsymbol{\theta} + \eta_{i,T-1} \quad (1.10) \\
S_{i,T} &= 1 \text{ if } S^*_{i,T} > 0, 0 \text{ otherwise,}
\end{aligned}$$

where the term $[\phi(.)/\Phi(.)]$ is called the inverse Mills ratio (IMR), ϕ is a standard normal density, and Φ is the standard normal cumulative distribution function; $\mathbf{z}_{i,T-1}$ and $\mathbf{y}_{s,T-1}$ are variable vectors that proxy the firms' expectations in period $T-1$ with regard to the decision to become an MNE. Exemplary estimates of the corresponding coefficients of these variables are depicted in table 1.6.[23] $S_{i,T} = 1$ means a firm chooses to change status from national to multinational, and $S^*_{i,T}$ is a latent variable that describes the propensity to invest. The error terms $\epsilon_{i,T}$ and $\eta_{i,T-1}$ are supposed to be bivariate, normally distributed with correlation ρ ($[\epsilon_{i,T}, \eta_{i,T-1}] \sim$ bivariate normal $[0, 0, 1, \sigma_\epsilon, \rho]$). The parameters in equation (1.9) are identified due to the non-linearity of the IMR, the time lags of the covariates in the probit, and the exclusion of indirect location specific variables from the main equation. Including the variable vectors $\mathbf{z}_{i,T-1}$ and $\mathbf{y}_{s,T-1}$ increases the sample periods to 1996-2000, 1996-1999,

[23]Post-investment growth rates of TFP O.P. 1, employment, average wages, and capital intensities are estimated using specification (2) of table 1.6. The growth of TFP O.P. 2 refers to specification (2), TFP L.P. to column (4), and labour productivity to specification (3).

and 1996-1998, respectively. Hence, vis-à-vis the OLS regressions of equation (1.8) Heckman's procedure causes a reduction in the number of observations, which occurs due to the unbalanced panel structure of the data. To arrive at comparable results, sample sizes of the OLS estimates were artificially reduced to match the observations of the Heckman (1978) estimator.

The coefficient β_1 of the $Switch_{i,T}$ dummy in (1.9) measures the *average treatment effect* (ATE), i.e. the expected impact of the switching decision on a randomly drawn firm. As opposed to the *average treatment effect on the treated* (ATT), the ATE also makes statements on units that would never be eligible for treatment. This problem can be reduced by excluding firms from the population that do not fit into a reasonable comparison group.[24] Through the exclusion of existing MNEs, switchers before and after the date of the regime change, observations smaller and younger than certain thresholds, and domestic holdings (Nace4 sectors 6523 and 7415) I adjust the estimation sample accordingly.

Table 1.7 reports annual growth rate premia in the years after switching. Cells in columns (1), (3), and (5) depict OLS estimates, columns (2), (4), and (6) include results using the Heckman (1978) estimator. Likelihood ratio (LR) tests on the independence of equations (1.9) and (1.10) allow to formally test for the occurrence of selectivity issues. All χ^2 statistics of the LR-tests are depicted beneath the respective numbers of observations.

The LR tests indicate that self-selection matters. Results for TFP growth in the first year after going multinational, labour productivity in the first and second period after becoming an MNE, average wages in the third year, and capital intensities for the whole sample period are significantly influenced by

[24]For a more elaborate discussion of this problem, see Wooldridge (2002).

the endogeneity of the switching decision. Moreover, corresponding parameter estimates without selection correction confirm the conjecture that OLS estimates are downward biased. In fact, most of the OLS results exhibit no significance at the 1%-10% level.

Average TFP growth in the first year after the regime change is significantly higher at newly founded MNEs than at national firms. Further significant growth differences with respect to TFP are found over a horizon of three years (Heckit model, $T + 3$). However, endogenity tests do not suggest to implement a selection correction and OLS results exhibit no significance in the three year period. A firm's productivity benefits from the decision to become a multinational with a growth premium of 5-7 percentage points immediately after switching, and with an extra annual growth rate of 4-5 percentage points over the three year horizon. Hence, these results provide evidence for the hypothesis that learning (through e.g. managerial and technological inputs from the foreign affiliate) increases the productivity at the parent operation.

Growth premia for capital intensities are significant for all years under consideration and 11-15 percentage points higher at newly founded MNEs than at national firms. This development is consistent with the hypothesis that German firms keep their more capital-intensive production stages at home and might also be the reason behind the relative expansion path of labour productivity (2-9 percentage points) and averages wages (1-4 percentage points). With a faster growing ratio of capital to labour, one would expect superior (labour) productivity growth rates and hence faster growing wages. Finally, employment growth, though superior before switching, exhibits no significantly higher annual rates at newly founded multinationals after the regime change. In other words, these results indicate that firms prepare for a forthcoming expansion to foreign markets by hiring additional employees, but these workers

Table 1.7: DIFFERENCES IN PERFORMANCE GROWTH, MNEs VS. NATIONALS t YEARS AFTER SWITCHING, ALL SECTORS

	$T+1^{a)}$		$T+2$		$T+3$	
	OLS	Heckit	OLS	Heckit	OLS	Heckit
Employment	*.005*[b)]	-*.035*	*.009*	-*.033*	-*.006*	-*.026*
	(.011)	(.022)	(.009)	(.024)	(.010)	(.023)
TFP O.P. 1	*.009*	.067	-*.009*	*.016*	*.013*	.051
	(.014)	(.025)	(.009)	(.020)	(.009)	(.018)
TFP O.P. 2	*.008*	.058	-*.010*	*.011*	*.013*	.048
	(.014)	(.026)	(.009)	(.021)	(.009)	(.019)
TFP L.P.	*.011*	.047	-*.003*	*.008*	.017	.038
	(.012)	(.025)	(.008)	(.019)	(.008)	(.021)
Labour productivity	*.009*	.085	*.011*	.077	.020	.047
	(.012)	(.025)	(.009)	(.019)	(.009)	(.026)
Average wage	.026	.059	.012	.038	*.007*	.036
	(.010)	(.022)	(.007)	(.015)	(.006)	(.011)
Capital/labour	-*.010*	.110	*.015*	.150	*.026*	.118
	(.023)	(.046)	(.019)	(.036)	(.020)	(.039)
N[c)] employment	32,375 (269)		20,631 (180)		11,576 (102)	
LR test[d)], $\chi^2_1 =$	2.55		1.98		0.89	
N TFP O.P. 1	32,269 (267)		20,530 (178)		11,498 (100)	
LR test, $\chi^2_1 =$	3.58*		1.15		2.05	
N TFP O.P. 2	32,267 (267)		20,527 (178)		11,497 (100)	
LR test, $\chi^2_1 =$	2.64*		0.79		1.69	
N TFP L.P.	32,269 (267)		20,530 (178)		11,498 (100)	
LR test, $\chi^2_1 =$	1.63		0.30		0.51	
N labour prod.	32,447 (267)		20,687 (177)		11,610 (101)	
LR test, $\chi^2_1 =$	5.78**		5.67**		0.89	
N average wage	32,330 (268)		20,592 (180)		11,556 (102)	
LR test, $\chi^2_1 =$	1.43		1.83		3.29*	
N capital/labour	32,352 (269)		20,617 (180)		11,565 (102)	
LR test, $\chi^2_1 =$	3.44*		6.31***		3.19*	

Source: USTAN and MiDi, Deutsche Bundesbank 1997-2001, own calculations.

[a)] Growth rates are measured as yearly averages assessed over the three years after switching ($T+1$-$T+3$).

[b)] Each cell includes the coefficient of the $Switch_{i,T}$ dummy in a regression that is either based on OLS or the Heckman (1978) endogenous treatment (Heckit) estimator. Standard errors are in parenthesis. If a parameter fails to be significant at the 10% level, it is set in italics.

[c)] Each N refers to the number of observations in the different performance regressions. The number of treated observations ($Switch_{i,T} = 1$) are set in parenthesis. Performance measures of switchers are evaluated between 1997 and 2001. The formation of new MNEs is observed between 1997 and 2000. All existing MNEs as well as switchers before and after the time of switching are removed from the estimation sample.

[d)] Likelihood Ratio tests of the hypothesis that the error terms of the probit and the treatment equations are uncorrelated are conducted.

then seem to meet the companies' personnel requirements in the time period after switching.

Summing up, the above analysis presents evidence that becoming an MNE increases post-investment performance with respect to productivity and average wages. Capital intensities evolve towards the use of capital, and switching does not affect firm size.

1.8 Conclusions

Chapter 1 investigated the extent to which performance attributes of multinational enterprises exceed those of purely national firms, both before and after they have switched from national to multinational activities. For that purpose a range of firm characteristics is evaluated. At the time of switching, newly founded MNEs exhibit performance premia of 10% (average wages) to 100% (firm size) compared to their national counterparts. Further regressions show, that multinationals already outperform national firms in the run up to investing abroad. Throughout this time period, the performance gap ranges from 91%-103% with respect to firm size and exhibits values between 21%-53% for the different productivity measures. Moreover, future multinationals pay on average 10%-14% higher wages, and capital intensities exceed those of national firms by 15%-25%. Two-sample Kolmogorov-Smirnov (KS) tests on the equality of performance distributions confirm the above findings. The tests clearly show that distribution functions of all firm characteristics for nationals lie to the left of those for switchers. With regard to ex-ante growth rates it turns out that only firm size exhibits higher annual rates in future MNEs relative to nationals. These differences are between 2-5 percentage points. Section 1.6 turns to a model of the decision to go multinational. I find that prior success, in terms of size, productivity, and a high portion of intangible assets in total assets increase the probability to become an MNE.

The use of Heckman's (1978) endogenous treatment model shows that (after switching) selectivity issues matter. I find evidence that TFP in the first year after going multinational, labour productivity in the first and second period after becoming an MNE, average wages in the third year, and capital intensities for the whole sample period are significantly influenced by the endogeneity of the switching decision. The dimension of ex-post growth rate differences between newly founded MNEs and domestic firms is 1-4 percentage points with respect to wages and 2-9 percentage points for the different productivity measures. The growth rate premia of capital labour ratios are between 11-15 percentage points per year.

These results confirm the view that international expansions of domestic firms are an important channel to raise overall competitiveness. The decision to become a multinational enterprise strengthens domestic operations. However, one has to take into account that the presented results refer to short run developments. The evaluation of performance measures at existing multinationals over a longer time horizon is beyond the scope of this analysis and may yield different results.

Appendix A

A.1 Construction of total factor productivity

As for the estimation of total factor productivity, I classified the Ustan data set into seven different branches (see table A.3). For each firm within a certain industry the following Cobb-Douglas production function is considered:

$$y_{i,t} = \beta_0 + \beta_1 l_{i,t} + \beta_2 k_{i,t} + \gamma_1 a_{i,t} + \gamma_2 r_i + \gamma_3 t_t + \nu_{i,t} + u_{i,t}. \tag{A.1}$$

Lower case letters indicate logarithmic values of the corresponding variables. $Y_{i,t}$ is the valued added of firm i at time t, $L_{i,t}$ and $K_{i,t}$ are its labour and capital inputs, $A_{i,t}$ is the firm age, r_i is a regional dummy that distinguishes East- and West-German locations, t_t is a linear time trend, $\nu_{i,t}$ is the part of productivity (unobservable for the researcher) that influences the firm's input decision, and $u_{i,t}$ includes both measurement error as well as unpredictable shocks to productivity.

Table A.1 exemplifies estimation results of the above equation for the sector *Wood, Chemicals and Others* during the period 1992 to 2001 using ordinary least squares (OLS), firm-specific unobserved effects (Within), the Olley and Pakes (O.P.) approach, and Levinsohn's and Petrin's (L.P.) method to control for endogeneity. A common feature of all estimation approaches is the assumption of constant, sector-specific production parameters over time, i.e. each firm active in the same industry produces with the same technology but with possibly different amounts of factor inputs. In columns (1), (2), and (4) the simple OLS estimator, which does not allow to treat $\nu_{i,t}$ and $u_{i,t}$ independently, is used. Columns (2) and (4) augment the first specification with the firms' age and investment as additional control variables. The within estimator of column (3) considers $\nu_{i,t}$ to vary over firms but to be constant over time. The results in columns (5)-(8) are based on semi-parametric estimation methods similar to the one proposed by Olley and Pakes (1996). Finally, column (9) is estimated using the Stata ado-file *levpet* (compare Levinsohn, Petrin and Poi 2003).[1]

The O.P. approach solves the endogeneity problem by expressing $\nu_{i,t}$ as a function

[1]Since the ado-file is very restrictive and does not allow to include other variables than capital and labour it mainly serves as an additional control specification.

of (contemporaneous) investment and capital. When applying this method, the following assumptions are made: (i) the inverted investment function can be written as $\nu_{i,t} = f(i_{i,t}, k_{i,t})$ ($i_{i,t}$ is the log of investment);[2] (ii) labour is the only variable factor, i.e. its demand is influenced by contemporaneous values of $\nu_{i,t}$; (iii) $k_{i,t}$ is a fixed variable influenced only by past values of the unobserved productivity shocks ($\nu_{i,t-1}$); (iv) $a_{i,t}$ is also fixed, but firms drop out of the sample randomly and hence past values of $\nu_{i,t}$ do no affect $a_{i,t}$.[3] Therefore, equation (A.1) changes to:

$$y_{i,t} = \beta_0 + \beta_1 l_{i,t} + \gamma_1 a_{i,t} + \gamma_2 r_i + \gamma_3 t_t + \phi_t(i_{i,t}, k_{i,t}) + u_{i,t}, \tag{A.2}$$

where $\phi_t = \beta_2 k_{i,t} + f(i_{i,t}, k_{i,t})$ is approximated by a 3rd order polynomial in log investment and log capital (including interaction terms). Equation (A.2) yields consistent estimates of β_0, β_1, γ_1, γ_2, and γ_3, while the coefficient of logarithmic capital β_2 is not identified. On this account, a second step is necessary to get consistent values of β_2. The second estimation equation is:

$$\begin{aligned} &y_{i,t+1} - \beta_0 - \beta_1 l_{i,t+1} - \gamma_1 a_{i,t+1} - \gamma_2 r_i - \gamma_3 t_{t+1} \\ &\quad = \beta_2 k_{i,t+1} + h(\phi_t - \beta_2 k_{i,t}) + \eta_{i,t+1} + u_{i,t+1}, \end{aligned} \tag{A.3}$$

where $\eta_{i,t+1}$ is the innovation in $\nu_{i,t+1}$ (defined as $\eta_{i,t+1} = \nu_{i,t+1} - E(\nu_{i,t+1}|\nu_{i,t})$), $\eta_{i,t+1}$ is mean independent of $k_{i,t+1}$ but possibly correlated with $l_{i,t+1}$, and $h(\cdot)$ is approximated by a 3rd order polynomial in $k_{i,t}$ and ϕ_t.

Estimation results of the O.P. approach are presented in columns (5) to (8).[4] None of the results in table A.1, columns (1), (2), and (4), account for the fact that ignoring $\nu_{i,t}$ causes an omitted variable bias. If the correlation between unobserved, firm-specific productivity shocks and the firm's factor demand is positive,

[2] For a derivation of the investment demand function see Olley and Pakes (1996).

[3] The last assumption is certainly questionable since firms leave markets when continuing the operation is not profitable. Unfortunately, it is not possible to correct estimations for this kind of attrition bias because firms in the USTAN data drop out of the sample if they do not draw a bill of exchange in a certain year.

[4] Table A.1 (columns (5) and (7)) also includes versions where $\nu_{i,t}$ is assumed to follow a random walk process ($\nu_{i,t+1} = \nu_{i,t} + \eta_{i,t+1}$). Equation (A.3) then reduces to:

$$y_{i,t+1} - \beta_0 - \beta_1 l_{i,t+1} - \gamma_1 a_{i,t+1} - \gamma_2 r_i - \gamma_3 t_{t+1} - \phi_t = \beta_2(k_{i,t+1} - k_{i,t}) + \eta_{i,t+1} + u_{i,t+1}.$$

OLS labour coefficients are upward biased. Turning to the O.P. and L.P. estimates should reduce much of the endogeneity problem. In fact, a comparison of columns (5)-(9) with columns (1), (2), and (4) reveals a decrease of the labour coefficient between eight percent in case of the different O.P. approaches and, with significantly more observations at hand, nineteen percent in case of Levinsohn's and Petrin's method. The within estimator, though suffering from the problem to model $\nu_{i,t}$ as constant over time, also provides evidence for a positive bias in the OLS estimates.

The capital coefficients turn out to be relatively low. Since firms with larger capital stocks are more likely to survive negative productivity shocks selecting only survivors should introduce negative correlation between the error term in the selected sample and the capital variable. As Olley and Pakes (1996) point out, this distortion could be reduced by including a selection correction term. Unfortunately, in estimations based on the USTAN data set an adjustment for this kind of bias is not possible because firms drop out of the sample if they do not draw a bill of exchange in a certain year. Hence, one needs to assume that panel attrition is caused by random drop outs of firms.

To focus on three productivity measures only, the results of columns (5), (7) and (9) are used to calculate total factor productivity. For these specifications my TFP measures are constructed as

$$TFP_{i,t} = exp(y_{i,t} - \beta_1 l_{i,t} - \beta_2 k_{i,t}[-\mathbf{c}_{i,t}\boldsymbol{\varphi}]), \tag{A.4}$$

where $\mathbf{c}_{i,t}$ are additional variables like firm age and other controls depending on the corresponding specification. In order to gain observations, I constructed out of sample predictions for firms where the investment or – in case of the L.P. estimates – intermediate input proxy was not available.

Table A.1: Estimation of Production Function Parameters, Sector: Wood, Chemicals and Others

	OLS	OLS age	Within	OLS inv.	O.P. 1 r.w.	O.P. 1 polyn.	O.P. 2 r.w.	O.P. 2 polyn.	L.P.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
ln employment	.874	.876	.579	.845	.803	.803	.805	.805	.622
	(.002)***	(.003)***	(.004)***	(.003)***	(.002)***	(.002)***	(.002)***	(.002)***	(.005)***
ln cap. stock	.131	.131	.067	.070	.083	.091	.084	.091	.071
	(.002)***	(.002)***	(.002)***	(.002)***	(.005)***	(.005)***	(.005)***	(.005)***	(.005)***
ln firm age	.	-.022	.149	-.016	.	.	-.009	-.009	.
		(.002)***	(.004)***	(.002)***			(.002)***	(.002)***	
ln investment	.	.	.	.082	.	.	.	.	.
				(.002)***					
region	.413	.420	.	.425	.431	.431	.431	.431	.
	(.007)***	(.008)***		(.009)***	(.006)***	(.006)***	(.007)***	(.007)***	
time trend	.016	.011	-.002	.011	.014	.014	.010	.010	.
	(.0006)***	(.0006)***	(.0004)***	(.0006)***	(.0006)***	(.0006)***	(.0006)***	(.0006)***	
constant	3.370	3.453	.	3.541	4.141	4.141	4.140	4.140	.
	(.010)***	(.011)***	(.020)***	(.011)***	(.035)***	(.035)***	(.040)***	(.040)***	
N 2nd step	.	.	.	.	45,995	45,995	37,629	37,629	.
N 1st step	85,175	65,804	66,312	53,902	69,103	69,103	53,902	53,902	87,400

Source: Ustan, Deutsche Bundesbank 1992-2001, own calculations.
Note: The dependent variable in columns (1) - (4) is the log of value added. The dependent variable in columns (5) and (7) is $y_{i,t+1} - \hat{\beta}_1 l_{i,t+1}[-\hat{\gamma}_1 a_{i,t+1}] - \hat{\beta}_0 - \hat{\gamma}_2 r_i - \hat{\gamma}_3 t_{t+1} - \hat{\phi}_{i,t}$, where $\hat{\phi}_{i,t}$ is approximated by a 3rd order polynomial in log investments and log capital from the 1st step estimation. The dependent variable in columns (6) and (8) is $y_{i,t+1} - \hat{\beta}_l l_{i,t+1}[-\hat{\gamma}_1 a_{i,t+1}] - \hat{\beta}_0 - \hat{\gamma}_2 r_{i,t} - \hat{\gamma}_3 t_{t+1}$ and (9) is estimated using the Stata ado-file levpet (compare Levinsohn, Petrin and Poi (2003)).

A.2 Currency conversion and deflation of foreign affiliate variables[5]

Becker, Ekholm, Jäckle and Muendler converted all economic data of foreign affiliates into euro (EUR) and deflated them. In BuBa's original MiDi data, all information on foreign affiliates is reported in German currency, using the exchange rate at the closing date of the foreign affiliate's balance sheet. The following deflation and currency conversion methods are applied to all (foreign affiliate) financial variables. (i) The market exchange rate on the end-of-month day closest to an affiliate's balance sheet closing date is used to convert the deutschmark (DEM) figures into local currency for every affiliate. This reverses the conversion applied to the questionnaires at the date of reporting. (ii) A deflation factor for every country deflates the foreign-currency financial figures to the December-1998 real value in local currency. (iii) For each country, the average of all end-of-month exchange rates vis-a-vis the DEM between January 1996 and December 2001 is used as a proxy for the purchasing power parity of foreign consumption baskets relative to the DEM. All deflated local-currency figures are converted back to DEM using this purchasing-power proxy. The resulting deutschmark figures are then converted into euro figures at the rate 1.95583 (the conversion rate at inception of the euro in 1999).

The foreign countries' CPIs (Consumer Price Indices from the IMF's International Financial Statistics, IFS) is used to deflate the figures. Whenever a country's CPI is not available from IFS but the main currency used in that country is issued in some other country, the CPI of the currency-issuing country is employed. The CPI deflation factors for all countries are rebased to unity at year-end 1998.

[5]Currency conversions and deflations were accomplished within the context of a broader research project on the impact of outward FDI on domestic labour markets. This section was first depicted in Becker, Ekholm, Jäckle and Muendler (2005b).

A.3 Summary statistics and sector definitions

Table A.2: Summary statistics, employment and capital stock

	Employment					
	33% pctl.	50% pctl.	67% pctl.	mean	obs.	overall obs.[c]
1993	23.00[a]	40.00	71.67	329.45	47,641	74,456
1994	23.33	40.33	73.33	331.6	49,089	75,021
1995	23.00	39.33	71.00	335.89	51,331	71,544
1996	22.67	39.00	70.33	343.83	50,840	69,423
1997	23.67	41.33	75.67	376.54	45,054	62,341
1998	26.00	47.33	90.33	475.63	35,072	48,194
1999	29.00	53.00	105.00	541.71	30,432	41,102
2000	31.67	60.33	120.33	595.57	27,343	36,207
2001	34.67	64.67	127.67	688.85	20,718	26,737
	Capital stock[b]					
	33% pctl.	50% pctl.	67% pctl.	mean	obs.	overall obs.
1993	178.946	423.116	942.301	3404.67	69,924	74,456
1994	180.469	432.547	986.070	3901.77	70,145	75,021
1995	184.351	442.832	1014.550	4078.39	66,913	71,544
1996	175.284	431.900	1010.687	4164.61	64,851	69,423
1997	176.316	455.497	1083.698	4688.06	58,103	62,341
1998	217.469	574.692	1405.201	5710.04	44,541	48,194
1999	259.164	714.522	1802.214	7024.70	37,798	41,102
2000	304.918	841.287	2160.430	7906.59	33,257	36,207
2001	355.587	942.101	2395.854	8041.65	24,601	26,737

Source: Ustan, Deutsche Bundesbank 1993-2001, own calculations.
[a] The tables depict summary statistics for the overall Ustan data set without any further adaptations.
[b] The capital stock is measured in thousands.
[c] Without missing values.

Table A.3: Aggregated sector definitions

1	Agriculture and mining
2	Food and textiles
3	Machinery and equipment
4	Wood, chemicals and others
5	Commerce
6	Finance and business
7	Other services

References

Arnold, Jens Matthias and Katrin Hussinger, "Export Behavior and Firm Productivity in German Manufacturing," *Review of World Economics*, 2005, *141* (2), 219–43.

—— **and** ——, "Export versus FDI in German Manufacturing: Firm Performance and Participation in International Markets," *ZEW Discussion Paper*, 2005, *No. 05-73.*

Becker, Sascha O., Karolina Ekholm, Robert Jäckle, and Marc-Andreas Muendler, "Location Choice and Employment Decisions: A Comparison of German and Swedish Multinationals," *Review of World Economics*, 2005, *141* (4), 693–731.

——**,** ——**,** ——**, and** ——, "Location Choice and Employment Decisions: A Comparison of German and Swedish Multinationals," *CESifo Working Paper*, January 2005, *1374. ssrn.com/abstract=648081.*

Bernard, Andrew and Bradford Jensen, "Exceptional exporter performance: cause, effect, or both?," *Journal of International Economics*, 1999, *47*, 1–25.

Blomström, Magnus, Gunnar Fors, and Robert E. Lipsey, "Foreign Direct Investment and Employment: Home Country Experience in the United States and Sweden," *Economic Journal*, November 1997, *107* (445), 1787–97.

Castellani, Davide and Giorgio Barba Navaretti, "Investments Abroad and Performance at Home: Evidence from Italian Multinationals," *CEPR Discussion Paper*, 2004, *4284.*

Caves, Richard E., *Multinational Enterprises and Economic Analysis*, 2d ed. ed., London: Cambridge University Press, 1996.

Clerides, Sofronis K., Saul Lach, and James R. Tybout, "Is Learning by Exporting Important? Micro-Dynamic Evidence from Colombia, Mexico, and Morocco," *The Quarterly Journal of Economics*, August 1998, *113* (3), 903–947.

Criscuolo, Chiara and Ralf Martin, "Multinationals, foreign ownership and US productivity leadership: Evidence from the UK," *mimeo*, April 2003.

Deutsche Bundesbank, "The Methodological Basis of the Deutsche Bundesbank's Corporate Balance Sheet Statistics," Monthly Report, 1998(10), 49-64.

Egger, Peter and Michael Pfaffermayr, "The Counterfactual to Investing Abroad: An Endogenous Treatment Approach of Foreign Affiliate Activity," *University of Innsbruck Working Papers in Economics*, 2003, *2.*

Girma, Sourafel and Richard Greenaway David Kneller, "Does Exporting Lead to Better Performance? A Microeconometric Analysis of Matched Firms.," *University if Nottingham Research Paper Series*, 2002, *09.*

___ **, Richard Kneller, and Mauro Pisu**, "Exports versus FDI: An empirical Test," *Review of World Economics*, 2005, *141* (2).

Heckman, James J., "Dummy Endogenous Variables in a Simultaneous Equation System," *Econometrica*, 1978, *46* (4), 931–60.

Helpman, Elhanan, Marc J. Melitz, and Stephen R. Yeaple, "Export versus FDI with Heterogeneous Firms," *The American Economic Review*, 2004, *94* (1), 300–316.

Levinsohn, James, Amil Petrin, and Brian P. Poi, "Production Function Estimation in Stata using Inputs to Control for Unobservables," *The Stata Journal*, 2003, *4* (2), 113–23.

___ **and** ___ , "Estimating Production Functions Using Inputs to Control for Unobservables," *Review of Economic Studies*, April 2003, *70* (2), 317–41.

Lipponer, Alexander, "A "New" Micro Database for German FDI," in Heinz Herrmann and Robert Lipsey, eds., *Foreign Direct Investment in the Real and Financial Sector of Industrial Countries*, Berlin: Springer, 2003, pp. 215–44.

Marin, Dalia, "A Nation of Poets and Thinkers—Less so with Eastern Enlargement? Austria and Germany," *CEPR Discussion Paper*, 2004, *4358.*

Markusen, James R., *Multinational Firms and the Theory of International Trade*, Cambridge, Massachusetts, London, England: The MIT Press, 2002.

Marschak, Jacob and William H. Jr. Andrews, "Random Simultaneous Equations and the Theory of Production," *Economatrica*, July-October 1944, *12* (3/4), 145–205.

Muendler, Marc-Andreas and Sascha O. Becker, "Margins of Multinational Labor Substitution," *IZA Discussion Paper No. 2131*, May 2006.

Navaretti, Giorgio Barba and Anthony J. Venables, *Multinational firms in the world economy*, Princeton and Oxford: Princeton University Press, 2004.

Olley, G. Steven and Ariel Pakes, "The Dynamics of Productivity in the Telecommunications Equipment Industry," *Econometrica*, November 1996, *64* (6), 1263–97.

Pfaffermayr, Michael, "Ownership Advantages, Foreign Production and Productivity: Evidence from Austrian Manufacturing Firms," *Review of Industrial Organization*, 1999, *15*, 37996.

Smirnov, Nikolai V., "Estimate of deviation between empirical distribution functions in two independent samples (in Russian)," *Bulletin Moscow University*, 1939, *2* (2), 3–16.

Wooldridge, Jeffrey M., *Econometric analysis of cross section and panel data*, Cambridge and London: MIT Press, 2002.

Chapter 2

The Impact of FDI on the Skill Structure in German Manufacturing

This chapter tests whether foreign direct investment (FDI) of German manufacturing multinationals (MNE) has raised domestic skill intensity between 1996 and 2001. Using a sample of 1,557 firms, the results show that foreign activities of German manufacturing MNEs carry higher average wages on the home market. I interpret this as evidence indicating that part of the skill upgrading in German manufacturing is associated with the rising job export to foreign locations. Other things equal, an increase in overall affiliate employment relative to domestic employment by 10 percentage points raises skill intensity at the parent firm by 0.1% to 0.3%. When distinguishing between different host regions, I find investment in industrialised countries consistent with the horizontal FDI motive, whereas investment in developing countries is driven by vertical production strategies. In the case of transition countries results are inconclusive.

2.1 Introduction[1]

In Germany (as in many other countries) an increase of multinational activities towards the end of the 20th century occurred. Between 1989 and 2001 the number of domestically owned multinational enterprises (MNE) rose from 6,762 to 8,857.[2] At the same time, German firms built up or acquired another 15,196 subsidiaries abroad, totalling 33,527 foreign affiliates in 2001. The workforce at these operations amounted to 1.720 million employees in 1989 and rose to 4.549 million workers in 2001. By the end of the corresponding time period, German firms produced 448.71 billion Euros' worth of goods abroad.

Trade theory suggests that outsourcing is associated with changes in relative labour demand. Horizontal FDI, i.e. replicating domestic operations at different locations, could either keep labour demand unaffected or shift it towards the more skilled. The effects of vertical FDI, on the other hand, depend on whether the host country is skill-abundant relative to the domestic country or vice versa. FDI flows to nations, which are abundant with low-skilled workers compared to Germany, may bring skill upgrading at the parent firm, whereas FDI to countries with a highly qualified workforce might be accompanied by a decrease in the skill level at home. On the firm level this translates to a changing share of the non-production wage bill and changing average wages.

Given those theoretical considerations, this chapter investigates whether the international diversification strategy of German manufacturing MNEs has influenced the domestic skill mix between 1996 and 2001. For that purpose, I employ three different foreign activity measures for German-headquartered multinationals: The ratios of (aggregated) foreign affiliate to domestic employment, output, and capital. Using these FDI proxies, augmented demand

[1]This chapter is an extended version of Jäckle (2006b).

[2]see Becker, Jäckle and Muendler (2005b).

functions for high-skilled labour are estimated.

The main finding is that foreign activities of German manufacturing MNEs carry higher average wages at domestic operations. I interpret this as evidence indicating that part of the skill upgrading in German manufacturing is associated with the international production strategy of German firms. Other things equal, a rise in overall affiliate employment relative to domestic employment by 10 percentage points is accompanied by an increase in the skill intensity at the parent firm by 0.1% to 0.3%. When distinguishing between different host regions, I find investment in industrialised countries consistent with the horizontal FDI motive, whereas investment in developing countries is clearly driven by vertical production strategies. In the case of transition countries results are inconclusive, a distinction between the two motives is not possible.

The remainder of this chapter is organised as follows: the starting point is a discussion of differences between horizontal and vertical investment strategies and their impact on the domestic skill structure, followed by a brief summary of the existing literature in section 2.3, the next section gives an overview of the data and provides several descriptives, then I look at specification issues and discuss the estimation results in section 2.5, and finally section 2.6 concludes the chapter.

2.2 Theoretical Considerations

The goal of this chapter is to investigate the effects of multinational activities on the skill structure at their domestic locations. Theoretical answers to this question are not clear cut. Instead, the outcomes are influenced by what motives lie behind a firm's decision to start up or acquire an affiliate in a certain region of the world. In this respect, the literature distinguishes between two types of FDI – *vertical* and *horizontal* FDI.[3] Since models based on one theory or the other come to different conclusions, the task to examine whether foreign investment causes skill upgrading or downgrading must be addressed empirically.

2.2.1 Vertical FDI

Cost saving efforts are the most important driving force toward companies accomplishing vertical FDI.[4] A firm decides to geographically fragment its production in separate stages, as a means to profit from differences in relative factor prices between the home and the host country. These benefits need to carry transport costs for the re-export of final or intermediate goods and fixed costs for starting up new production facilities abroad. Markusen (2002) summarises in his book that "for vertical firms, location advantages arise when trade costs are low, stages of production differ in factor intensities, and countries differ significantly in relative factor endowments." Therefore, the effect of vertical FDI on the skill structure at home mostly depends on whether the domestic country is abundant in skilled labour relative to a large proportion of unskilled workers in the foreign country, or vice versa. In the first case, which is most likely for investment in low-income countries, one would expect that knowledge based assets are still produced at the parent firm, whereas final and intermediate production stages are accomplished at the affiliate operation. Hence, FDI flows to countries, which are relatively abundant with low-skilled workers, may come along with skill upgrading at the parent firm. In the lat-

[3] Markusen, Venables, Konan and Zhang (1996) and Carr, Markusen and Maskus (2001) combine the two models to the so called "knowledge-capital" model. Through its three defining assumptions *fragmentation, skilled-labour intensity, and jointness* the model allows for vertical and horizontal activities in one common framework.

[4] Among the first authors who described vertical multinationals were Helpman (1984, 1985), and Helpman and Krugman (1985).

ter case, which mostly refers to investment in high-income countries, theory suggests a decrease in the skill level at home.

For Germany, a country which is relatively abundant with skilled labour, it is often argued that investment in low wage countries (especially those located in Central and Eastern Europe (CEE)) may put pressure on domestic (low-skilled) wages. Hence, according to the vertical FDI theory it seems straightforward that an investment strategy, which fragments production such that labour intensive stages are located in CEE countries, Asian-Pacific countries, or other developing countries, increases the skill intensity in German manufacturing.

Marin (2004), on the other hand, finds evidence that Austrian and German companies take advantage of cheap high-skilled workers in Central and Eastern Europe. She argues that German firms locate headquarter activities like research and development (R&D) in low-wage countries. One may refer to this as some kind of inverted vertical FDI, i.e. instead of taking advantage of the abundant low skilled labour German firms (additionally) replace parts of their knowledge based assets at home with cheap (high-skilled) labour in transition economies. In this case, one would expect that, ceteris paribus, the skill level at the parent firm decreases.

Reaffirming the above, it is a priori not clear whether the vertical division of the production process upgrades or downgrades the skill intensity in Germany. In the end, it depends on which of the two opposing forces – the relocation of low- or high-skilled labour – is stronger. If the first effect dominates one would expect vertical FDI to positively influence the skill structure. If, on the other hand, indirect vertical fragmentation turns out to be most important, there might be no effect or even a downgrade of the skill intensity.

2.2.2 Horizontal FDI

If a firm conducts horizontal FDI, aspirations for better market access and the proximity-concentration trade-off play a decisive role. Companies invest abroad if the costs of accessing new markets and transportation are higher than the expenditures of starting up a new firm and the loss due to a reduction of scale economies, when producing the same good across different markets. In this case, Markusen (2002) states that "for horizontal firms, location advantages arise when the host country market is large and when trade costs (broadly defined) are moderate to high." Head and Ries (2002) bring an additional point into the discussion. They argue that horizontal FDI can be both a replication of all activities at different locations or only a replication of final goods production (in their terminology "*branching*"), where upstream stages of production, like design and marketing, are located at home. In the latter case, at least part of the firms' (skill intensive) knowledge based assets stay at the parent location.

If production in foreign countries is independent from local factors (*replication*) there exists only an indirect effect on the domestic firm through possible changes in the scale of the parent operation. These changes may occur if production at foreign affiliates substitutes for exports to the according markets. Head and Ries (2002) show that a scale-decrease in the domestic location may be accompanied be either skill-upgrading or downgrading.[5] Yet, the important point at this juncture is that there exists no effect which is independent of changes in the scale of the parent operation. Hence, in an empirical specification where one controls for the size of the parent firm (compare section 2.5)

[5]In the extreme case where the demand for high skilled employees is completely independent of the output produced, i.e. knowledge based inputs require only a certain amount of high-skilled workers (independent of output), the skill intensity at the domestic operation increases if domestic production is reduced.

the expected impact of outward FDI on the skill structure with respect to the theory of horizontal replication should not be measured as significantly different from zero.

In the *branching* case, where knowledge capital produced at home serves as a common input for production in subsidiaries all over the world and horizontal FDI is meant to build up production capacities for final goods only, controlling for the size of the domestic firm in a regression approach does not reduce the effect on the parent skill level to zero. The argument at this juncture is that in a case of accessing new markets (where no goods were exported to so far) the overall increase in worldwide production corresponds to an expansion of knowledge-based input factors at home, whereas domestic output stays constant. Therefore, the skill intensity at the parent firm will be enhanced. If, on the other hand, *branching-investment* acts as a substitute for exports the scale of domestic production will be reduced. Given that the reduction at home is fully compensated by new capacities abroad, worldwide productions requires still the same input of knowledge capital. Therefore, even after controlling for the size of the domestic operation one should find a positive effect of *branching* on the skill level at the parent firm.[6]

For Germany, one might distinguish between vertical and horizontal FDI by discriminating between the different foreign locations German firms choose. High income region, especially in Western Europe and North America, may serve as locations where mainly horizontal FDI (replication and branching) is conducted. In the case of developing and transition countries, however, the distinction is less clear. The motive for locating affiliates in these regions might

[6]In the branching case, the skill increase at the parent operation cannot be controlled by including the size of the parent firm since, in contrast to the replication scenario, the (additional) demand for knowledge capital comes through the "affiliate-channel".

be driven by both horizontal and vertical aspects. Since developing and transition countries are mostly abundant in low-skilled labour, a substantial part of investment made there may replace relatively expansive unskilled employees at home, thus leading to a higher skill intensity at the domestic location. To the contrary, many of these countries already serve as important markets for final goods. This is true, for instance, in the emerging markets of China and India, but also many of the transition countries in Central and Eastern Europe.

2.3 Related Literature

A number of recent studies examine the consequences of the international division of production for labour markets in industrialised countries. Among the first empirical papers written in this field are Feenstra and Hanson (1996, 1999). Their focus lies on the influence of globally integrated production on the observed wage divergence in the US between 1979 and 1990. The authors find evidence that outsourcing activities (defined as imports of intermediate inputs from their own and foreign affiliates) of American firms contributed significantly to the relative increase of high skilled wages in the considered time period. Instead of defining broad outsourcing measures, Slaughter (2000) solely concentrates on the question whether production transfers through foreign direct investment has contributed to the increased demand for high skilled workers in US manufacturing. In the descriptive part of his paper he shows that between 1977 and 1994 US multinationals extended foreign activities relative to domestic ones. However, his regression results suggest that in a hypothetical situation without increased MNE transfers the observed skill upgrading would have been almost the same. The results of Feenstra and Hanson (1999) along with those of Slaughter (2000) can be seen as evidence that for the US the wage gap is mainly effected by trade at arm's length, sub-contracting, and licensing instead of direct investment. For Sweden, Hansson (2001) shows that transfers of production stages from Swedish MNEs to non-OECD countries positively affect relative wages of skilled workers. In a study of Japan, Head and Ries (2002) investigate the impact of overseas employment of Japanese enterprises on the skill structure at the domestic location. They use data that is aggregated on the sector level as well as micro-data at the level of the parent firm. Their results suggest that expanding the work force in low-income countries brings an upgrade of the skill level at parent operations in Japan.

Geishecker and Görg (2004) are among the first to ask who the "Winners and Losers" from outsourcing in Germany are. They estimate wage equations separately for three (low, medium, and high) skill groups and augment the different specifications using foreign activity measures as defined by Feenstra and Hanson (1996, 1999). Their results suggest that workers in the lowest skill category lose from outsourcing and that – at least in some specifications – high-skilled workers gain through higher wages. In a paper based on aggregated industry data, Geishecker (2006) finds a significant negative effect of outsourcing towards Central and Eastern Europe on the relative demand for manual workers in German manufacturing. In another study of Germany, Jäckle (2006a) looks amongst other performance measures at average wages of firms both prior to and after they have gone multinational (switching). He shows that compared to non-switchers average wages have already been superior in the run up to become a MNE, and secondly, that after switching average wages grow faster at newly founded multinationals than at purely domestic firms.

2.4 Data and Descriptives

In the study at hand, data from the Deutsche Bundesbank's MiDi database at the level of German manufacturing parents and their (aggregated) foreign affiliate activities between 1996 and 2001 is used. From the MiDi data set information on ownership-weighted foreign employment, fixed assets, and turnover is gathered. All financial variables referring to affiliate operations are converted into euro and deflated to real values at year end 1998 employing a purchasing-power related method (for further details see section 1.3 and appendix A.2). Information on parent specific variables is obtained from the BuBa Ustan data set (see Deutsche Bundesbank 1998), which was string matched by name to companies in the MiDi data base. The variables extracted from Ustan are employment, turnover, fixed assets, overall labour costs, and intermediate input goods.[7] All financial figures except intermediate input goods are deflated to 1998 real values using the German CPI (from the IMFs International Financial Statistics). Intermediate input goods are converted to real values using the the intermediate input goods deflator from the OECD's Main Economic Indicators. The value added is constructed as the difference between real turnover and real intermediate input goods.

Both the MiDi as well as the Ustan data set are available in the form of an unbalanced panel. Over the whole sample period, a total of 1,557 different manufacturing firms were matched. Table B.1 in the appendix includes industry definitions and reports the panel attrition. In table 2.1 the development of the data sets between 1996 and 2001 is depicted. The first line reports the overall number of firms in German manufacturing with at least one foreign af-

[7]Turnover variables in both the MiDi and Ustan data sets do not distinguish between within-MNE shipments of final goods and final sales of parent and affiliate firms. Nonetheless, these variables are the best available proxies to parent firms' production at home and affiliates' output abroad.

filiate. The second row includes the corresponding number of foreign affiliates. These figures show that both the number of parent firms (+8.5%) as well as the number of affiliates (+12.4%) rose in the time period under consideration. A comparison with lines three and four, which include the total number of matched firms in each year, allows an evaluation of the matching algorithm. The merge process yields a matching quote between 40 percent in 1996 and 29 percent in 2001 for parent firms and a coverage between 54 percent (1996) and 42 percent (2001) for their foreign affiliates.

Table 2.1: Number of FDI firms and foreign affiliates in German manufacturing

	1996	1997	1998	1999	2000	2001
	(1)	(2)	(3)	(4)	(5)	(6)
Parent firms total	2,596	2,710	2,782	2,782	2,849	2,817
Foreign affiliates total	10,403	10,935	11,432	11,375	11,828	11,689
Parent firms matched	1,034	1,047	993	969	929	811
Foreign affiliates matched	5,639	5,787	5,860	5,506	5,558	4,949

Source: Ustan and MiDi, Deutsche Bundesbank 1996-2001, own calculations.

Table 2.2 depicts the worldwide (domestic and foreign) usage of employment and capital of German manufacturing firms between 1996 and 2001. The table additionally focuses on the output produced with these inputs and on average labour costs incurred per employee. While affiliate numbers are aggregated values based on ownership-weighted firm-level observations from the MiDi data set (MiDi includes the overall population of German firms conducting outward FDI), sector-wide values are based on aggregated numbers for the whole manufacturing sector (including also non-MNEs) provided by the national account series of the German Federal Statistical Office. Table 2.3 presents the regional pattern of the MNEs' aggregated foreign activities for three broad country groups. The groupings are: Industrialised countries (*IND*), transition countries (*TRANS*), and developing countries (*DEV*) (for definitions see table B.3 in the appendix). All tables also include the corresponding figures for the matched sample of MiDi and Ustan firms, used in the econometric analysis of section 2.5.

Table 2.2: EMPLOYMENT, TURNOVER, AND CAPITAL OF GERMAN MANUFACTURING

	1996	1997	1998	1999	2000	2001
	(1)	(2)	(3)	(4)	(5)	(6)
Total observations						
Foreign affiliates						
Employment	1,292.616	1,343.883	1,531.898	1,504.524	1,562.200	1,538.254
Fixed assets	78.248	77.422	96.532	100.297	124.638	129.285
Turnover	299.733	319.484	420.904	461.512	473.793	449.038
Parent sector						
Employment	11,194	10,903	10,814	10,652	10,591	10,417
Fixed assets	1,472.648	1,463.241	1,473.63	1,476.097	1,435.436	1,435.443
Avg. wages	35,491	35,419	35,728	36,134	36,844	36,788
Turnover	1,368.982	1,399.725	1,447.08	1,480.97	1,536.808	1,540.342
Matched sample						
Foreign affiliates						
Employment	836.398	786.018	794.413	757.393	780.564	741.339
Fixed assets	60.083	49.292	50.700	49.860	53.837	50.100
Turnover	219.683	198.565	198.983	189.167	204.036	191.468
Parent firms						
Employment	1,376.887	1,316.960	1,272.875	1,196.742	1,205.114	1,082.388
Fixed assets	127.128	137.454	145.456	144.761	163.804	156.093
Avg. wages	45,330	45,602	48,248	49,140	52,269	47,654
Turnover	292.789	307.974	318.963	304.930	333.855	310.724

Source: Ustan and MiDi, Deutsche Bundesbank 1996-2001, own calculations, and Federal Statistical Office Germany, Fachserie 18/Reihe 1.4. Foreign variables are ownership-weighted. All financial variables are deflated to unity at year end 1998 and, with the exception of average labour costs, measured in billions of Euros. Employment figures are in thousands.

Excluding self-employed persons, in 2001 the absolute number of employees in the German manufacturing sector (10.417 million) was 6.9% lower than in 1996 (11.194 million). At the same time, overall employment at foreign affiliates increased by 19% from 1.293 million workers in 1996 to about 1.538 million employees in 2001. The decline of parent employment in the matched sample (-21.4%) virtually equals the decrease in sample size between 1996 and 2001 (-21.6%) (see table 2.1). Hence, one may conclude that the employment reduction at manufacturing MNEs in Germany was less strong than in the whole sector. Assuming that absolute home market employment at manufacturing MNEs stayed constant, relative multinational-wide employment shifted towards foreign operations. These numbers therefore provide descriptive evidence, confirming the public opinion, that an increasing number of jobs at foreign affiliates substitute for parent employment in Germany.[8]

While the number of workers in German manufacturing declined, the figures in table 2.2 show that – at the same time – real average wages increased.[9] Looking at the matched sample, the development towards higher average wages per firm is even stronger. However, it is unfortunate that there is a relatively strong drop in average wages in 2001, which does not correspond to the sector-wide development, and might be due to the decreasing sample size in the last period. In sum, the increase in average wages by 3.7% together with the declining employment (-6.9%) in German manufacturing could be seen as an initial piece of descriptive evidence for skill-upgrading during the time period under consideration.

Apart from variables related to the firms' workforce, table 2.2 also focuses

[8]Using a translog cost function approach, Becker et al. (2005a) find that affiliate employment tends to be a substitute for parent employment. In a follow up study, Muendler and Becker (2006) show how multinational labour demand responds to wage differentials at the extensive margin, when a multinational enterprise expands, and at the intensive margin, when an MNE operates existing affiliates. They derive conditions to infer elasticities of labor substitution at both margins, controlling for location selectivity. Their results show that with every percentage increase in German wages, German MNEs allocate 1,600 manufacturing jobs to Eastern Europe at the extensive margin and 3,900 jobs overall.

[9]The skill proxy used in the regressions of section 2.5 is the average wage paid by each firm. It is constructed using the USTAN variables wage bill and employment (see appendix). To construct comparable average wages for the overall sector (from the national account series of the German Federal Statistical Office), I used sector-wide labour costs (including social security contributions paid by the employer) and divided those by the number of workers (without self-employed).

on capital usage and output. It shows that domestic firms reduced fixed assets (-2.5%), while output simultaneously rose by 12.5%. At their foreign affiliates, German manufacturing MNEs strongly expanded both the use of capital (65.2%) and the absolute value of output (49.8%). Since affiliate output expansion was proportionally larger, the multinational-wide distribution of turnover shifted towards affiliates. Furthermore, when looking at the capital distribution, it becomes clear that investment abroad must have been relatively larger than investment at home.

When looking at regional FDI patterns (see table 2.3), in 2001 industrialised countries are still the most important host region. At that time they accounted for 58.1% (64% in 1996) of affiliate employment, 85.9% (83.9%) of affiliate fixed assets, and 84.1% (85.1%) of the output German manufacturing firms produce abroad. In Germany, most of the public attention focuses on outward FDI to transition countries located in Central and Eastern Europe (CEE). In terms of employment numbers this country group raised its relative importance. German manufacturing MNEs increased their foreign employment share in transition countries from 10.7% in 1996 to 17.5% in 2001. Furthermore, with respect to capital and output the corresponding figures amount to 2.5% and 2.8% in 1996 and 2.8% and accordingly 3.1% in 2001. Finally, in developing host countries the relative share of foreign activities in 2001 was slightly lower than in 1996 (employment and fixed assets) or kept almost constant during the sample period (turnover).

Table 2.3: Foreign employment, turnover, and capital of German manufacturing MNEs, differentiated by world regions

	1996	1997	1998	1999	2000	2001
	(1)	(2)	(3)	(4)	(5)	(6)
Total observations						
Industrialised countries						
Employment	827.727	825.508	957.416	914.149	909.902	893.121
Fixed assets	65.643	64.191	80.542	82.215	106.233	111.098
Turnover	255.155	269.034	362.389	401.026	403.075	377.692
Transition countries						
Employment	137.904	171.250	208.797	224.931	261.306	269.212
Fixed assets	1.987	2.295	3.041	3.258	3.606	3.579
Turnover	7.374	9.709	12.414	13.922	15.566	13.869
Developing countries						
Employment	326.985	347.125	365.685	365.444	390.991	375.922
Fixed assets	10.617	10.937	12.949	14.825	14.800	14.608
Turnover	37.205	40.741	46.101	46.564	55.151	57.477
Matched sample						
Industrialised countries						
Employment	541.002	488.567	484.489	444.784	433.839	415.411
Fixed assets	50.871	40.176	40.501	38.593	42.176	39.510
Turnover	185.107	164.583	164.215	153.703	160.938	150.571
Transition countries						
Employment	76.881	91.220	106.542	111.508	136.257	134.454
Fixed assets	1.389	1.678	2.209	2.231	2.384	2.197
Turnover	5.630	6.864	8.897	9.801	10.818	8.778
Developing countries						
Employment	218.516	206.231	203.382	201.101	210.468	191.473
Fixed assets	7.823	7.438	7.990	9.035	9.277	8.392
Turnover	28.946	27.117	25.871	25.663	32.279	32.119

Source: Ustan and MiDi, Deutsche Bundesbank 1996-2001, own calculations. The variables are ownership-weighted. All financial variables are measured in billions EUR and deflated to unity at year end 1998. Employment figures are in thousands.

2.5 Specification Issues and Econometric Results

The figures in section 2.4 provide descriptive evidence that increasing foreign activities of German manufacturing firms in the time period between 1996 and 2001 were associated with higher average wages and, at the same time, decreasing employment numbers. However, to infer that these trends reflect within-firm shifts of labour demand from low-skilled employees towards the more high-skilled, one needs to turn to regression analysis.

2.5.1 Specification

The existing literature tries to understand, what explains changes in the skill structure, in a translog cost function framework. Throughout the different studies, the authors assume that capital inputs are a quasi-fixed factor and that firms/industries minimise their costs with respect to low- and high-skilled workers. Berman, Bound and Griliches (1994) use this approach to derive the corresponding (production and non-production) share equations and investigate whether labour-saving technological change shifts demand away from low-skilled employees. Feenstra and Hanson (1996, 1999) augment the Berman et al. (1994) specification and additionally include outsourcing proxies in their equations. Slaughter (2000) uses the cost function approach to explore what the impact of FDI on the skill structure in US manufacturing is. All of the above studies have in common that they use data on the industry or sector level. Unlike those, Head and Ries (2002) are the first who look at FDI-induced skill changes at the firm level.

In this chapter, I follow in their footsteps and look at the firms' demand equations for high skilled labour:[10]

$$S^h_{j,t} = \beta_0 + \beta_1 \log(\frac{w^h_{j,t}}{w^l_{j,t}}) + \beta_k \log \frac{K_{j,t}}{Y_{j,t}} + \beta_Y \log Y_{j,t} + \beta_m MNE_{j,t} + u_{j,t}, \quad (2.1)$$

where $S^h_{j,t}$ is the share of skilled workers at firm j in year t, $w^h_{j,t}$ and $w^l_{j,t}$ are the respective input prices for low and high skilled labour, $K_{j,t}$ are fixed assets, $Y_{j,t}$ is the real value added output, $MNE_{j,t}$ is a variable that measures the multinational activity of firm j in year t, and $u_{j,t}$ is an unobserved error term.

The parameter β_k accounts for the impact of capital intensity on the skill structure. It is positive if capital and skilled employees exhibit a complementary relationship. The value added regressors β_Y accounts for the size of the firm. Most importantly, the variable $MNE_{j,t}$ controls for the variation in $S^h_{j,t}$ that is due to changes of the firms' FDI-activities. I follow Slaughter (2000) and employ three different activity measures: affiliate employment, turnover, and fixed assets. I compute these measures by building the sum over all affiliates that belong to the same parent firm. To avoid double counting, if one affiliate is owned by more than one German parent, each variable is weighted with its parent firm's ownership share.[11] Using these absolute measures to account for foreign activities does not allow to control for general equilibrium shocks, that might change domestic and foreign performance. A worldwide recession, for example, could hit both parent and affiliate employment. Therefore, I construct the regressors $MNE_{j,t}$ as the ratio of (aggregated) foreign affiliate to domestic employment, output, and capital.[12] Employing the two non-labour

[10] See section B.2 for a formal derivation of the translog cost function approach.

[11] The ownership variable also allows the restriction of the estimation sample in section 2.5 to majority owned affiliates. The use of the restricted sample did not significantly change the regression outcome. Modified estimation results are available on request.

[12] In contrast to the analysis at hand, Head and Ries (2002) use the share of a MNE's total work force that is located offshore to quantify the firm's international activities. For comparison reasons, I also constructed the regressors $MNE_{j,t}$ as affiliate activity divided by worldwide (domestic + foreign) activity. The use of these different measures did not alter

measures turnover and fixed assets allows to test, whether investing abroad accounts for more than just transferring jobs. Since FDI might additionally cause affiliate capital and output expansion, using these measures enables me to test for their impact on the skill mix in German manufacturing.

The dependent variable in my regressions is the log of average wages (normalised by average sectoral wages) paid by each parent firm. The mean wage is meant to proxy the skill intensity of production and is computed as real labour costs over employment. The normalisation controls for productivity gains that are common to all workers. In the appendix (B.1) conditions, under which average wages are a good approximation for the high-skill share ($S^h_{j,t}$) of a firm's workforce, are derived. While the rough estimates in appendix B.1 suggest that the log of average wages over average low-skilled wages can serve as an indicator for the skill intensity, two problems arise at this point. First, since the USTAN data set does not include any information about the skill composition of the firms' workforce, I have to rely on average yearly industry-wages from the German Statistical Office to normalise average wages.[13] Second, as Head and Ries (2002) point out, average wages on the firm level might be an indicator for efficiency wages. Thus, as far as the efficiency mark-ups between firms differ, part of the variation in average wages might be explained independently of the employees' skill levels. Burdett and Mortensen (1998) show that wage differentials and firm size are positively correlated, i.e. larger firms pay efficiency wages to reduce employment fluctuations. Therefore, including firm size as an explanatory variable should help to control for the efficiency-part

results in any important way.

[13] In other specifications, I also used average yearly wages of low-skilled workers from the German Socio-Economic Panel (GSOEP) to normalise average labour costs. I classify the workers' educational attainment as low-skilled if they have either no school degree at all, no school degree plus vocational training, or a lower school degree without vocational training. However, the different denominator of the dependent variable did not alter results in any important way. Estimation outcomes, using these modifications, are available on request.

of the wage variation. Additionally, the inclusion of firm fixed-effects in most specifications might also account for this kind of variation. Finally, I follow Berman et al. (1994) and Head and Ries (2002), assuming that quality-adjusted high- and low-skilled labour does not vary over firms. Under this assumption, relative input prices of the two skill groups are constant, and time dummies are used to capture yearly changes in the wage rate for all firms.

It is thus the case that equation (2.1) takes the following estimatable form:

$$log(\frac{\bar{w}_{j,t}}{\bar{w}^{l}_{j,t}}) = \beta_0 + \delta year_t + \beta_k \log\frac{K_{j,t}}{Y_{j,t}} + \beta_Y \log Y_{j,t} + \beta_m MNE_{j,t} + c_j + \epsilon_{j,t}, \quad (2.2)$$

where $\bar{w}_{j,t}$ is the average wage paid by firm j in year t, $\bar{w}^{l}_{j,t}$ is the average low skilled wage, $year_t$ are yearly time dummies, c_j is an unobserved firm specific factor, and $\epsilon_{j,t}$ is an error term, which is assumed to be mean independent of the explanatory variables given c_j. The decisive parameter in equation (2.2) is β_m. It measures the impact of FDI on the skill mix in German manufacturing. Firm fixed effects (c_j) are included to account for unobservable variables, which are correlated with the key variable $MNE_{j,t}$ – the standard omitted variable bias problem – and do not vary over time. Good examples in this respect are latent management skills and objectives. To compare, I also include pooled estimation results (without latent constant effects) in tables 2.4 and 2.5.

2.5.2 Results

Table 2.4 provides estimation results for equation (2.2). Columns (1) through (3) report specifications using pooled regressions, columns (4) through (6) include parameter estimates for the fixed effects model.

For manufacturing firms in Germany, equation (2.1) explains between 7%

and 17% (7% and 26%) of the variance in (demeaned) log normalised average wages in the pooled sample (fixed effects model). Including relative affiliate employment instead of affiliate fixed assets or output increases the (within) R^2 from a mere 7% to 17% (7% to 26%). This can be seen as the first piece of evidence, demonstrating that job creation at foreign subsidiaries better explains the variation in domestic skill composition than the expansion of affiliate output and capital.

Corresponding to results obtained by other researchers, higher levels of output significantly increase average wages across all specifications. Burdett and Mortensen (1998) argue that wage differentials and firm size depend on each other, i.e. larger firms pay efficiency wages to reduce employment fluctuations. Hence, estimation results indicate that the output variable controls for this type of scale effects. Furthermore, a positive wage-scale relationship could indicate that successful firms (in terms of output) pass on part of their gains to the workforce (e.g. in the form of higher bonuses) and herewith increase average wages.

The pooled regression model exhibits significant, positive effects of the capital to value added ratio, i.e. an expansion of capital usage is accompanied by higher average wages. However, estimated coefficients in the fixed effects model differ greatly from those without unobserved constant factors. As opposed to the pooled regression model, the inclusion of c_j reveals a significant, negative influence of capital on the skill intensity. This means that estimated parameters in the pooled specifications might be severely upward biased. The substitutional relationship between capital and skill in German manufacturing stands in contrast to Slaughter's US sector-level study, which finds the two input factors to be complementary. Head and Ries (2002), on the other hand, show that for Japanese MNEs – similar to German firms – greater capital investment is negatively correlated with average wages.

Table 2.4: Multinational activities and average wages in German manufacturing, results

	pooled (1)	pooled (2)	pooled (3)	within (4)	within (5)	within (6)
Activity measure	employmt.	turnover	fixed ass.	employmt.	turnover	fixed ass.
MNE activity	.009 (.002)***	.0007 (.0005)	.001 (.0002)***	.026 (.004)***	.0003 (.0003)	-.0001 (.001)
Log capital/val.-add.	.024 (.007)***	.042 (.008)***	.043 (.008)***	-.038 (.011)***	-.031 (.012)***	-.031 (.012)***
Log value-added	.048 (.003)***	.041 (.004)***	.041 (.004)***	.162 (.035)***	.125 (.041)***	.109 (.041)***
Dummy 1997	.0005 (.011)	.004 (.012)	.006 (.012)	-.002 (.006)	.007 (.007)	.007 (.007)
Dummy 1998	.021 (.012)*	.021 (.012)*	.023 (.012)*	.006 (.007)	.019 (.008)**	.019 (.008)**
Dummy 1999	.010 (.012)	.016 (.012)	.017 (.012)	.004 (.007)	.017 (.008)**	.019 (.008)**
Dummy 2000	.023 (.012)*	.037 (.014)***	.042 (.014)***	.023 (.008)***	.043 (.011)***	.046 (.011)***
Dummy 2001	.024 (.011)**	.031 (.012)**	.031 (.012)**	.023 (.008)***	.045 (.010)***	.046 (.010)***
Constant	-.488 (.063)***	-.347 (.070)***	-.340 (.070)***	.	.	.
N	5458	5185	5210	5458	5185	5210
N manufac. firms	1480	1394	1404	1480	1394	1404
R^2	.166	.077	.074	.261	.074	.066
F-test time dummies	1.853*	2.52**	2.761**	3.681***	5.164***	5.497***

Source: USTAN and MiDi, Deutsche Bundesbank 1996-2001, own calculations. Pooled and firm fixed effects estimation. Standard errors are in parenthesis, where * denotes significance at the 10%, ** at the 5%, and *** at the 1% percent level. Standard errors are estimated using the Huber/White/sandwich adjustment.

Across all specifications, F-tests of joint significance of coefficients on the time dummies refer to regressions including a full set of yearly binary variables. Positive, significant parameter estimates on year dummies show that skill upgrading in German manufacturing cannot be fully explained by capital, output, and MNE activities. The coefficients are largest (and highly significant) in the years 2000 and 2001, which might be a response to the higher panel attrition at the end of the sample period (compare also table 2.1 in section 2.4 and table B.1 in the appendix to this chapter).

Turning to the effects of multinational activities on the skill intensity, the central result is that FDI, especially when proxied by foreign employment, is one of the driving forces behind skill-upgrading in German manufacturing. The figures in table 2.4 indicate, that (i) in the pooled regression approach multinational activity measures, reflecting affiliate employment and capital, are significant and positive, and (ii) when including constant unobserved effects, only the coefficient on affiliate jobs positively affects skill intensity at the parent firm. Since the dependent variable exhibits a logarithmic form, β_m can be interpreted as growth rate of normalised average wages. Other things equal, a rise in affiliate employment relative to domestic employment by 10 percentage points is associated with an increase in the skill intensity at the parent firm between 0.09% in the pooled regression and 0.26% in the fixed effects model.

Table 2.5 includes results, where multinational activity measures are separated into three broad country groups. The groupings are: Industrialised countries (*IND*), transition countries (*TRANS*), and developing countries (*DEV*) (for definitions see table B.3 in the appendix). Estimation results with respect to domestic capital investment and value added lie within the range of the corresponding parameters depicted in table 2.4. Again, higher levels of output

Table 2.5: Multinational activities and average wages in German manufacturing, results differentiated by world regions

	pooled	pooled	pooled	within	within	within
	(1)	(2)	(3)	(4)	(5)	(6)
Activity measure	employmt.	turnover	fixed ass.	employmt.	turnover	fixed ass.
MNE activity *IND*	.010	.0003	.006	.025	.0001	.0003
	(.004)**	(.0002)*	(.003)*	(.009)***	(.0001)	(.0009)
MNE activity *TRANS*	.006	.066	-.010	.010	.064	-.004
	(.006)	(.021)***	(.006)	(.006)	(.011)***	(.008)
MNE activity *DEV*	.010	.272	.003	.029	.240	.0003
	(.006)*	(.072)***	(.004)	(.004)***	(.202)	(.009)
Log capital/val.-add.	.024	.028	.044	-.037	-.037	-.031
	(.007)***	(.007)***	(.008)***	(.011)***	(.011)***	(.012)***
Log value added	.048	.040	.041	.161	.141	.109
	(.003)***	(.004)***	(.004)***	(.036)***	(.041)***	(.041)***
Dummy 1997	.0004	.004	.007	-.001	.007	.007
	(.011)	(.012)	(.012)	(.006)	(.007)	(.007)
Dummy 1998	.021	.019	.023	.007	.016	.019
	(.012)*	(.012)	(.012)*	(.007)	(.008)*	(.008)**
Dummy 1999	.010	.015	.018	.006	.013	.019
	(.012)	(.012)	(.012)	(.007)	(.009)	(.008)**
Dummy 2000	.023	.039	.042	.024	.036	.046
	(.012)*	(.013)**	(.014)***	(.008)***	(.011)***	(.011)***
Dummy 2001	.024	.031	.031	.025	.038	.046
	(.011)**	(.012)**	(.012)**	(.008)***	(.011)***	(.010)***
Constant	-.488	-.342	-.347	.	.	.
	(.065)***	(.069)***	(.069)***			
N	5458	5185	5210	5458	5185	5210
N manufac. firms	1480	1394	1404	1480	1394	1404
R^2	.167	.141	.075	.267	.111	.066
F-test time dummies	1.878*	2.357**	2.824**	3.96***	4.07***	5.557***

Source: Ustan and MiDi, Deutsche Bundesbank 1996-2001, own calculations. Pooled and firm fixed effects estimation. Standard errors are in parenthesis, where * denotes significance at the 10%, ** at the 5%, and *** at the 1% percent level. Standard errors are estimated using the Huber/White/sandwich adjustment.

contribute significantly to skill upgrading, and a larger capital to value added ratio comes along with lower skill levels at parent firms (fixed effects model). The same holds true for yearly time dummies. As in the specification with only one common foreign activity measure, I find significantly positive coefficients on many of the binary time variables and F-tests of joint significance refer to regressions including a full set of year dummies.

As for MNE activities in *industrialised* countries, the pooled regression approach suggests that higher affiliate employment, turnover, and capital brings rising average wages in German manufacturing. In the fixed effects model, only increasing affiliate employment is associated with higher skill intensities at the parent operation. Assuming that developed regions attract FDI of the horizontal type, estimation results suggest that most of the subsidiaries located in these countries replicate final good production at home. At the same time, upstream activities might stay on the domestic market, and a positive effect on the skill level at parent firms may occur. Therefore, at this juncture, evidence for the *branching-hypothesis* of Head and Ries (2002) (see section 2.2.2) is found.

When turning to *transition* countries, for both the pooled and the fixed effects model affiliate output is the only foreign activity measure that significantly affects domestic skill intensity. Rising affiliate turnover is associated with increasing average wages at home. These findings could indicate that (i) FDI in these countries is of the horizontal branching-type and mainly driven by market access motives, or/and (ii) German manufacturing firms vertically divide their production process, locating final goods production in transition countries, and re-export finished goods.

Finally, when focusing on *developing* countries, skill-upgrading in German manufacturing can be explained by job and output transfers in the case of pooled OLS, and when including unobserved constant effects only affiliate employment positively affects domestic wages. In the latter case, a 10 percentage points increase in the workforce in developing countries is associated with an 0.29% skill increase at the parent operation. Since developing countries are relatively abundant with low-skilled labour compared to Germany, my findings confirm that German manufacturing MNEs exploit relative factor price differences (= vertical FDI) when investing at these locations.

2.6 Conclusions

This chapter analysed the impact of the international diversification strategy of German manufacturing MNEs on the domestic skill mix between 1996 and 2001. The descriptive figures in section 2.4, which show increasing foreign activities, decreasing home employment, and rising average wages, suggest a shift in labour demand towards the more skilled. To confirm that these trends actually reflect within-firm changes of the skill structure of manufacturing firms, a translog cost function approach is employed and demand functions for high-skilled labour are estimated.

The main finding of this paper is that foreign activities of German manufacturing MNEs are positively correlated with higher average wages at domestic operations. I interpret this as evidence that part of the skill upgrading in German manufacturing is associated with the rising job export to foreign locations. Other things equal, an increase in overall affiliate employment relative to domestic employment by 10 percentage points is accompanied by an increase in the skill intensity at the parent firm by 0.1% to 0.3%. When distinguishing between different host regions, I find investment in industrialised countries consistent with the horizontal FDI motive, whereas investment in developing countries is driven by vertical production strategies. In the case of transition countries results are inconclusive, a distinction between the two motives is not possible.

Appendix B

B.1 Are average wages a good proxy for the skill intensity?

The study at hand proxies the skill intensity in German manufacturing using average wages paid by each firm on the home market. It is constructed as the wage bill divided by employment, where both variables are available in the USTAN data set. By splitting up the overall wage bill into the sum of earnings of high- and low-skilled employees one obtains:[1]

$$\bar{w}_j = \frac{\sum_{i=1}^{N_j} w_{ij}}{N_j} = \frac{\sum_{i=1}^{N_j^l} w_{ij}^l + \sum_{i=N_j^l+1}^{N_j} w_{ij}^h}{N_j^l + N_j^h} = \frac{\bar{w}_j^l N_j^l + \bar{w}_j^h N_j^h}{N_j^l + N_j^h}, \tag{B.1}$$

where $\bar{w}_j$ is the average wage paid by firm j, N_j is the overall number of workers employed by firm j, N_j^l and N_j^h are low- and high-skilled employment, and $\bar{w}_j^l$ and $\bar{w}_j^h$ denote average wages for low- and high-skilled employees at firm j, respectively. Some further transformation of equation (B.1) yields:

$$\bar{w}_j = \frac{\bar{w}_j^l}{1 - \frac{\bar{w}_j^h N_j^h}{\bar{w}_j^h N_j^h + \bar{w}_j^l N_j^l} \times \frac{\bar{w}_j^h - \bar{w}_j^l}{\bar{w}_j^h}} = \frac{\bar{w}_j^l}{1 - I_j \times P_j}, \tag{B.2}$$

where I_j is the skill intensity at firm j, and P_j is wage premium of skilled over unskilled employees. Dividing both sides of equation (B.2) by $\bar{w}_j^l$ and taking logs results in:

$$\log(\frac{\bar{w}_j}{\bar{w}_j^l}) = -log(1 - I_j \times P_j) \approx I_j \times P_j. \tag{B.3}$$

The second equality holds only for small values of the product of skill intensities and wage premia. Equation B.3 states that the log of average payments to a firm's domestic workforce over the average low-skilled wage is roughly proportional to the skill intensity at company j, given the skill premium on wages is constant in the period under consideration.

According to calculations using the German Socio-Economic Panel (GSOEP)

[1] The following derivations are based on Head and Ries (2002).

the premium for higher education as opposed to basic education (P_j) in Germany between 1996 and 2001 remains roughly constant at 41%.[2] At the same time, the average skill-intensity in Germany is between 33% and 41%.[3] Therefore, the respective values of $-log(1 - I_j \times P_j)$ lie in between 0.15 and 0.18 and the approximated values of $I_j \times P_j$ are 0.14 and 0.17, respectively.

B.2 The transcendental logarithmic (translog) cost function

Assuming that capital can be treated as a quasi-fixed factor, the (short-run) translog cost function can be written in the following form:[4]

$$\begin{aligned}
log(C) &= log(\alpha_0) + \alpha_Y log(Y) + \alpha_K log(K) + \sum_i \alpha_i log(w_i) + \\
&+ 0.5\beta_{YY} log(Y)^2 + 0.5\beta_{KK} log(K)^2 + \\
&+ 0.5 \sum_i \sum_j \beta_{ij} log(w_i) \times log(w_j) \\
&+ \sum_i \gamma_{iY} log(w_i) \times log(Y) + \sum_i \gamma_{iK} log(w_i) \times log(K), \qquad \text{(B.4)}
\end{aligned}$$

where $i \neq j; i, j = 1, ..., n$, Y is the value added, K is the capital stock, and $w_{i/j}$ are input prices. Applying equation (B.4) to the two (variable) input case (low and

[2] I classify the workers' educational attainment as low-skilled if they have either no educational degree at all, no school degree plus vocational training, or a lower school degree without vocational training. High-skill employment refers to persons with a high-school degree plus additional vocational training, higher technical college, or a university degree. To calculate average wages, fulltime gross earnings (incl. 13th month salary, vacation and Christmas bonus) of the respective skill groups are employed. Since I made use of the samples A-F of the GSOEP a weighting scheme is used to overcome the problem of different sampling probabilities when inferring average values of the target population.

[3] A rough estimate of the above numbers can be obtained using the overall wage sum and the wage sum of unskilled employees from the GSOEP. Again, all observations are weighted according to their sample probabilities.

[4] This following derivations are based on Berman, Bound and Griliches (1993) and Chung (1994).

high skilled labour) results in the following equation:

$$\begin{aligned} log(C) &= log(\alpha_0) + \alpha_Y log(Y) + \alpha_K log(K) + \alpha_h log(w^h) + \alpha_l log(w^l) + \\ &+ 0.5\beta_{YY} log(Y)^2 + 0.5\beta_{KK} log(K)^2 + \\ &+ 0.5\beta_{ll} log(w^l)^2 + 0.5\beta_{hh} log(w^h)^2 + \\ &+ \beta_{hl} log(w^h) \times log(w^l) + \\ &+ \gamma_{lY} log(w^l) \times log(Y) + \gamma_{hY} log(w^h) \times log(Y) + \\ &+ \gamma_{lK} log(w^l) \times log(K) + \gamma_{hK} log(w^h) \times log(K). \end{aligned} \tag{B.5}$$

Shepard's (1953) Lemma suggests that the first order differentiation of the cost function with respect to an input price yields the cost-minimizing demand function for the corresponding input factor. Therefore, differentiating equation (B.5) with respect to $log(w^h)$ yields:

$$\begin{aligned} S^h &= \frac{N^h \times w^h}{C} = \frac{\partial C}{\partial w^h} \times \frac{w^h}{C} = \frac{\partial log(C)}{\partial log(w^h)} = \\ &= \alpha_h + \beta_{hh} log(w^h) + \beta_{hl} log(w^l) + \gamma_{hY} log(Y) + \gamma_{hK} log(K), \end{aligned} \tag{B.6}$$

where S^h is the cost share and N^h is the number of skilled employees.

The cost function is constrained to be homogenous of degree one in input prices if the following restrictions are satisfied: (i) $\alpha_l + \alpha_h = 1$; (ii) $\sum_{i=1}^{2} \beta_{i,j} = \sum_{j=1}^{2} \beta_{j,i} = 0 \Longleftrightarrow \beta_{hl} = -\beta_{hh}$; (iii) $\gamma_{lY} + \gamma_{hY} = 0$; (iv) $\gamma_{lK} + \gamma_{hK} = 0$. Hence, equation (B.6) changes to:

$$S^h = \alpha_h + \beta_{hh} log(\frac{w^h}{w^l}) + \gamma_{hY} log(Y) + \gamma_{hK} log(K). \tag{B.7}$$

Finally, under the assumption that $\gamma_{hK} = -\gamma_{hY}$ (i.e. constant returns to scale; see Berman et al. 1993) and by following the usual practice of adding a foreign activity measure (MNE) and including Y as a proxy for firm size (see Slaughter 2000) I arrive (by slightly abusing notation in (B.7)) at the equivalent of equation (2.1) in section 2.5.1:

$$S^h = \beta_0 + \beta_1 \log(\frac{w^h}{w^l}) + \beta_k \log\frac{K}{Y} + \beta_Y \log Y + \beta_m MNE. \tag{B.8}$$

B.3 Summary statistics, industry and regional definitions

Table B.1: Industry definitions and panel attrition

	1996	1997	1998	1999	2000	2001	total
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Food products and beverages	53	54	43	40	34	34	258
Tobacco products	$\leq 3^{a)}$	≤ 3	≤ 3	≤ 3	≤ 3	≤ 3	10
Textiles	42	42	42	39	35	35	235
Wearing apparel; dressing of furn.	38	30	28	31	27	22	176
Tanning and dressing of leather	8	10	11	9	7	5	50
Wood and cork (no furniture)	10	14	13	12	9	10	68
Pulp, paper and paper products	19	20	19	16	15	12	101
Publishing and printing	13	13	16	18	18	13	91
Coke, petroleum, nuclear fuel	5	5	5	4	≤ 3	≤ 3	23
Chemicals	103	110	98	93	94	80	578
Rubber and plastic	60	65	62	65	60	50	362
Other non-metallic	43	45	40	34	29	27	218
Basic metals	45	39	31	32	39	37	223
Fabricated metal	101	105	97	93	90	78	564
Machinery and equipment	265	266	261	247	241	213	1493
Office machinery and computer	8	7	7	≤ 3	4	≤ 3	31
Electrical machinery	58	56	55	62	57	48	336
Communication equipment	17	16	17	23	24	22	119
Medical and precision instr.	64	60	63	66	62	47	362
Motor vehicles, trailers	42	46	46	47	50	45	276
Other transport equipment	5	5	6	5	5	4	30
Furniture	33	36	30	26	24	23	172
Recycling	0	≤ 3	≤ 3	≤ 3	≤ 3	≤ 3	7
total	1034	1047	993	969	929	811	$5783^{b)}$

Source: Ustan and MiDi, Deutsche Bundesbank 1996-2001, own calculations.
[a)] Data protection guidelines of the Deutsche Bundesbank oblige researchers to hide figures where the number of observed firms is smaller than three.
[b)] 5,783 observations from 1,557 unique firms between 1996 and 2001.

Table B.2: SUMMARY STATISTICS

	N	Mean	Std. dev.	10% pctl.	90 % pctl.
	(1)	(2)	(3)	(4)	(5)
Dependent variable					
Log average wage normalised	5670	.366	.318	.074	.634
Regressors					
MNE activities employment	5606	1.620	11.572	.023	1.570
MNE activities turnover	5312	7.108	196.442	.033	1.050
MNE activities fixed assets	5430	.573	6.686	.010	.961
Log capital/value added	5596	-.776	.891	-1.826	.229
Log value added	5596	17.421	1.415	15.862	19.185
Regressors regional specification					
MNE activities *IND*					
Employment	5606	.937	7.807	0	.854
Turnover	5312	6.798	194.036	0	.937
Fixed assets	5430	.279	2.360	0	.565
MNE activities *TRANS*					
Employment	5606	.355	3.743	0	.333
Turnover	5312	.040	.766	0	.055
Fixed assets	5430	.142	1.541	0	.259
MNE activities *DEV*					
Employment	5606	.328	4.191	0	.194
Turnover	5312	.270	9.232	0	.085
Fixed assets	5430	.151	2.983	0	.190

Source: USTAN and MIDI, Deutsche Bundesbank 1996-2001, own calculations. All summary statistics are on firm-year level. Regressors are ownership weighted. The dependent variable is normalized using industry-year wages form the German Statistical Office. Foreign activity variables are measured relative to domestic activities.

Table B.3: Regional definitions

IND	Western European countries (EU 15 plus Norway and Switzerland) and Overseas Industrialised countries (Canada, Japan, USA, Australia, New Zealand, as well as Iceland and Greenland)
TRANS	Central and Eastern European countries (accession countries and candidates for EU membership)
DEV	Developing countries (Asia-Pacific, Hong Kong, South Korea, Singapore, Taiwan, China, Mongolia, and North Korea; Russia and Central Asian economies) and other developing countries (South Asia (India/Pakistan), Africa, Latin America, the Middle East; including dominions of Western European countries and the United States)

References

Becker, Sascha O., Karolina Ekholm, Robert Jäckle, and Marc-Andreas Muendler, "Location Choice and Employment Decisions: A Comparison of German and Swedish Multinationals," *Review of World Economics*, 2005, *141* (4), 693–731.

——, **Robert Jäckle, and Marc-Andreas Muendler**, "Kehren deutsche Firmen ihrer Heimat den Rücken? Ausländische Direktinvestitionen deutscher Unternehmen," *Ifo Schnelldienst*, January 2005, *58* (1), 23–33.

Berman, Eli, John Bound, and Zvi Griliches, "Changes in the Demand for Skilled Labor within U.S. Manufacturing: Evidence from the Annual Survey of Manufacturers," *NBER Working Paper Series*, January 1993, (4255).

——, ——, **and** ——, "Changes in the Demand for Skilled Labor within U.S. Manufacturing: Evidence from the Annual Survey of Manufacturers," *The Quarterly Journal of Economics*, May 1994, *109* (2), 367–397.

Burdett, Kenneth and Dale T. Mortensen, "Wage Differentials, Employer Size, and Unemployment," *International Economic Review*, 1998, *39* (2), 257–273.

Carr, David L., James R. Markusen, and Keith E. Maskus, "Estimating the Knowledge-Capital Model of the Multinational Enterprise," *American Economic Review*, June 2001, *91* (3), 693–708.

Chung, Jae Wan, *Utility and Production Functions*, Cambridge: Blackwell Publishers, 1994.

Deutsche Bundesbank, "The Methodological Basis of the Deutsche Bundesbank's Corporate Balance Sheet Statistics," Monthly Report, 1998(10), 49-64.

Feenstra, Robert C. and Gordon H. Hanson, "Globalization, Outsourcing, and Wage Inequality," *American Economic Review*, 1996, *86* (2), 240–45.

—— **and** ——, "The Impact of Outsourcing and High-Technology Capital on Wages: Estimates for the United States, 1979-1990," *Quarterly Journal of Economics*, August 1999, *114* (3), 907–40.

Geishecker, Ingo, "Does Outsourcing to Central and Eastern Europe really threaten manual workers' jobs in Germany?," *The World Economy*, 2006, *29* (5), 559–583.

——— **and Holger Görg**, "Winners and Losers: Fragmentation, Trade and Wages Revisited," *IZA Discussion Paper*, January 2004, *982.*

Hansson, Pär, "Skill Upgrading and Production Transfer within Swedish Multinationals in the 1990s," *FIEF Trade Union Institute for Economic Research Stockholm Working Paper*, January 2001, *166.*

Head, Keith and John Ries, "Offshore Production and Skill Upgrading by Japanese Manufacturing Firms," *Journal of International Economics*, October 2002, *58* (1), 81–105.

Helpman, Elhanan, "A Simple Theory of International Trade with Multinational Corporations," *Journal of Political Economy*, March 1984, *92*, 451–71.

——— , "Multinational Corporations and Trade Structure," *Review of Economic Studies*, 1985, *52*, 443–58.

——— **and Paul Krugman**, *Market Structure and Foreign Trade*, Cambridge: MIT Press, 1985.

Jäckle, Robert, "Going Multinational: What are the effects on home market performance?," *Bundesbank Discussion Paper Series 1, No 03*, 2006.

——— , "The Impact of FDI on the Skill Structure in German Manufacturing," *Applied Economics Quarterly Supplement*, 2006, (52), 55–78.

Marin, Dalia, "A Nation of Poets and Thinkers—Less so with Eastern Enlargement? Austria and Germany," *CEPR Discussion Paper*, 2004, *4358.*

Markusen, James R., *Multinational Firms and the Theory of International Trade*, Cambridge, Massachusetts, London, England: The MIT Press, 2002.

——— **, Anthony J. Venables, Denisee E. Konan, and Kevin H. Zhang**, "A unified treatment of horizontal direct investment, vertical direct investment, and the pattern of trade in goods and services," *NBER Working Paper*, 1996, *5696.*

Muendler, Marc-Andreas and Sascha O. Becker, "Margins of Multinational Labor Substitution," *IZA Discussion Paper No. 2131*, May 2006.

Shepard, Ronald W., *Cost and production functions*, Princeton: Princeton University Press, 1953.

Slaughter, Matthew J., "Production Transfer within Multinational Enterprises and American Wages," *Journal of International Economics*, April 2000, *50* (2), 449–72.

Chapter 3

Health and Wages Panel data estimates considering selection and endogeneity

This paper investigates the effects of health on wages by controlling for a number of problems: first, the unobservable genetic endowment may cause an omitted variable bias; second, using a self-reported health variable could induce measurement error; third, the issue of reverse causality arises; and fourth, panel attrition driven by the endogenous decision to participate in the labour market may result in inconsistent estimation. By using recently developed methods, I control for all of the above issues in one framework. The results show that good health raises wages for both women and men. I find the health variable to suffer from measurement error. For men, applying OLS or 2SLS, instead of methods accounting for selection and individual heterogeneity, causes an upward bias in the health coefficient. Selection tests indicate panel attrition to generate biased estimates in the male sample, while for females no selection correction is required.

3.1 Introduction

Whether there exists a measurable interrelation between health and wages is an important question in both labour and health economics. There are two reasons which establish a link between the state of health and wages. First, health as part of one's human capital may affect labour market productivity and hence wages. Second, as Grossman (2001) points out, if marginal benefits of investment in health increase with the salary, health should rise with wages and the issue of reverse causality comes up. However, a number of further challenges arise. To start with, as self-reported health satisfaction is used for estimation, it is not possible to assess one's actual health status accurately and measurement error could be a source of bias. Another shortcoming that is unappreciated in most earlier studies of this kind is sample selection. Since labour market participation is endogenous – with one reason for selection being the health status – applying methods without selection corrections may result in inconsistent estimation. Finally, an issue particulary relevant in the health context is individual heterogeneity. The reasonable presumption that genetic endowment is correlated with health calls for panel data techniques to account for the well known omitted variable bias.

In an attempt to control for all of these problems in one framework, I utilise recently developed estimation methods proposed by Wooldridge (1995) and Semykina and Wooldridge (2005). In the first paper, Wooldridge develops new straightforward techniques to test and correct for sample selection in fixed effects models. His method is easier to implement and more flexible than other models in the literature as it does not demand any known distribution of the error terms in the main equation, and allows them to be time heteroscedastic and serially correlated in an unspecified way. In an application to female labour supply, Dustmann and Rochina-Barrachina (2000) compare Wooldridge's (1995) estimator to the methods proposed by Kyriazidou (1997) and Rochina-Barrachina (1999). Kyriazidou's (1997) estimator is semi-parametric and matches observations with the same selection effect in two periods. By taking the difference between any two years one gets rid of both individual heterogeneity and sample selection. A crucial point is the "conditional exchangeability" assumption, implying that the idiosyncratic errors are homoscedastic over time conditional on the covariates and unobserved effects in both equations. While Kyriazidou (1997) does not impose distributional assumptions on the selection term, Rochina-Barrachina (1999) parameterises this effect and assumes joint normality of the error terms in the probit and main equation. Her method does not rely on the "conditional exchangeability" assumption. Dustmann and Rochina-Barrachina (2000) show how to expand the three estimators to account for the problems of non-strict exogeneity and measurement error. Similarly, Semykina and Wooldridge (2005) enhance Wooldridge's (1995) estimator and demonstrate how to test and control for sample selection in a fixed effects model with endogeneity. Again, their approach allows for time heteroscedasticity and autocorrelation in the error terms in both equations.

Turning to the literature concerned with the impact of health on wages, there are several papers worth mentioning. To start with, Lee (1982) suggests an econometric model that accounts for the simultaneous effects of health and wages in a structural multi-equation system, based on a generalised version of the Heckman (1978) treatment model. Using a male sample of US citizens, he finds that health and wages are strongly interrelated; that is the wage rate positively affects health and vice versa. Haveman, Wolfe, Kreider and Stone (1994) estimate a multiple equation system for working time, wages, and health, employing generalised methods of moments techniques. In their male sample for the US they show that poor health affects wages negatively. Contoyannis and Rice (2001) study the impact of self-assessed general and psychological health on wages using the British Household Panel Survey. They apply fixed effects and random effects instrumental variable estimators and conclude that reduced psychological health decreases male wages, while positive self-assessed health increases hourly wages for women. In a recent paper, Gambin (2005) investigates the relationship between health and wages for 14 European countries and finds that for men, self-reported health has a greater effect than for females, while in the case of chronic diseases the opposite holds true.

In this chapter data from the German Socio-Economic Panel (GSOEP) is used to estimate reduced-form wage equations for women and men augmented by a variable measuring health satisfaction. I follow Wooldridge (1995) and Semykina and Wooldridge (2005) in an attempt to account for the problems of unobserved heterogeneity, sample selection, and endogeneity. A number of tests provide evidence that for the male sample selection corrections are indicated, while for women no selection problems occur. The results show that good health raises wages. For females an increase in health satisfaction by 10% enhances (hourly) wages approximately by 0.14 to 0.47 percent. In the male sample the increase of the wage rate ranges from about 0.09 to 0.88 percent. The health variable is found to suffer from measurement error. For men, employing pooled OLS or 2SLS, instead of methods accounting for selection and individual heterogeneity, is accompanied by an upward bias in the health coefficient.

The remainder of this chapter is organised as follows: the starting point is a discussion of specification issues and resulting problems; that is followed by an detailed overview of the different estimation methods in section 3.3; the next part provides summary statistics of the data; then, in section 3.5, I look at estimation and test results; and, finally, section 3.6 concludes the chapter.

3.2 Model Specification and Resulting Problems

In order to improve our understanding of how health affects wages, a simple model is presented. In this model, the only input factor is the quantity of effective labour L_t a firm uses to produce Y_t at time t. The production function of a firm is determined by the function $Y_t = F(L_t)$, and the amount of effective labour can be written as

$$L_t = \sum_{i=1}^{n} p_i(s_i, a_{i,t}, h_{i,t}) \times l_{i,t}, \tag{3.1}$$

where $l_{i,t}$ is the actual labour supply per employee i, and $p_i(\cdot)$ is a unknown function that determines the effectiveness of $l_{i,t}$. The efficiency of an individual's working hours depends on the (maximum) years of schooling s_i, age $a_{i,t}$, and her/his state of health $h_{i,t}$. In what follows, I refer to the first two variables as the human capital part of $p_i(\cdot)$ and to the latter part as health effect.

If workers are paid according to their marginal product the log wage of each employee can be written as

$$\log w_{i,t} = log[\frac{dF}{dL_t} \times \frac{\partial L_t}{\partial l_{i,t}}] = \log F_{L_t} + \log p_i(s_i, a_{i,t}, h_{i,t}). \tag{3.2}$$

This implies that log wages can be decomposed into the term $\log F_{L_t}$, which depends on supply and demand factors on the firm level, and a human capital and health effect, respectively, that varies on the level of the employee. In order to approximate the first part, I use yearly averages of job-seekers and notified vacancies on the level of the federal states in Germany,[1] as well as four different categories for the firm size. To find a plausible functional form for the human capital part of the term $\log p_i(\cdot)$, a specification of the variables $a_{i,t}$ and s_i similar to the one proposed by Mincer (1958 and 1974) is assumed. Finally, to cover the health status, a function of a self-assessed health measure is included, which asks individuals for a description of their current satisfaction with health.[2]

The following parameterization captures the above model:

$$\log w_{i,t} = \mathbf{b}_{s,t}\boldsymbol{\alpha} + \mathbf{f}_{i,t}\boldsymbol{\beta} + \mathbf{a}_{i,t}\boldsymbol{\gamma} + \theta s_i + \delta f(h_{i,t}) + error, \tag{3.3}$$

where $\mathbf{b}_{s,t}$ is a vector that approximates supply and demand forces on the (federal) state level s, $\mathbf{f}_{i,t}$ are dummy variables capturing different firm sizes, $\mathbf{a}_{i,t}$ is the vector of a 3rd order polynomial of $a_{i,t}$, s_i are years of schooling or training, $f(h_{i,t})$ is a function of the health variable, and $(\boldsymbol{\alpha}', \boldsymbol{\beta}', \boldsymbol{\gamma}', \theta, \delta)'$ is the corresponding parameter vector.

The Health Effect. There are a number of important links that connect the state of health and earnings. First, health as part of one's human capital affects labour market productivity and hence wages. Second, in the theoretical work of Grossman (2001), health is defined as an endogenous capital stock, which determines the amount of time one can spend in producing monetary income. Since average hours worked deviate substantially among individuals – with one reason for the difference being the health status – (the log of real) hourly wages

[1]The corresponding figures are extracted from "Arbeitsstatistik 2005 - Jahreszahlen", provided by the Federal Employment Agency, Nuremberg.

[2]The health variable is categorial, ranging form zero to ten. It is transformed using the following function: $f(h_{i,t}) = \log(h_{i,t} + \sqrt{(h_{i,t}^2 + 1)})$.

rather than monthly earnings are analyzed.[3] Third, in Grossman's (2001) model the rate of return to (gross) investment in health equals the additional availability of healthy time, evaluated at the hourly wage rate. This means that health should rise with wages as the marginal benefits of health investment increase with the wage rate, implying that $h_{i,t}$ is *simultaneously* determined along with $w_{i,t}$.

When estimating the $(\boldsymbol{\alpha}', \boldsymbol{\beta}', \boldsymbol{\gamma}', \theta, \delta)$ in equation (3.3) a number of further problems arise. To start with, *measurement error* can be an important source of bias when trying to explain wages by employing self-reported questions about health satisfaction. An example of an objective health measure would be a physician's diagnosis of a persons's biological state of health. However, in the absence of such a variable it is likely that δ will be biased towards zero. Another problem arises due to the non-availability of a random sample from the population. In this study, I am interested in the effect of health on the labour market productivity of *all* persons. So, taking into account only the working population induces a sample *selection problem*. In this context, a bias results from the fact that individuals endogenously decide to participate in the labour market. Since it is likely that some of the factors determining participation also affect health, the selection process might lead to inconsistent estimation. A further problem is the possible appearance of an *omitted variable bias*. In this respect one could think of the genetic endowment of a person. If somebody is genetical 'well' equipped she/he might at the same time be healthier and draw a higher salary, so that the health coefficient is upward biased. Finally, as has been noted by Contoyannis, Jones and Rice (2004) and Halliday and Burns (2005) it is likely that the state of health follows a per-

[3]This specification also suits equation (3.2) well since the derivative of $F(L_t)$ with respect to the actual working time, $l_{i,t}$, suggests utilising hourly wages as dependent variable in equation (3.3).

sistent stochastic process. The literature describes two sources of persistence: individual heterogeneity and state dependence. The first one exists due to the (unobserved) degree to which a person is able to cope with individual health shocks (such as hard attacks, accidents, etc.). State dependence, as the second source of persistence, means that an individual's ability to deal with health shocks depends on her/his (former) health status.

The major focus of this study is to control for simultaneity, measurement error, omitted variables, and selection in one common framework. Unfortunately, the methods proposed in section 3.3 do not allow to fully cover the dynamics in the state health. Persistence working through the (unobserved) individual ability to cope with health problems can be controlled for by including unobserved effects. Dynamic effects due to the state dependence of the health status, on the other hand, necessitate to include an (unknown) number of lagged health variables. Yet, the estimation of a 'complete' model identifying the above sources of endogeneity plus the full dynamics of health is beyond the scope of this study. Therefore, a parsimonious specification including only contemporaneous values of health satisfaction is employed. Non-inclusion of lagged health variables, however, leaves a source of endogeneity in the model which is controlled by applying an instrumental variable approach that uses lagged values of variables related to former health shocks (number of doctor visits in the last three months, number of days off from work due to illness last year).

The Human Capital Part. As mentioned before, the human capital part of $p_i(\cdot)$ is approximated using a Mincer-like specification. He suggests using a model, where log wages are linear in the years of schooling, and linear and quadratic in the years of labour market experience. In an empirical application using the GSOEP, Romeu Gordo (2006) finds evidence for the existence of a positive relationship between unemployment and health satisfaction. On this account, I decided to employ a specification that includes unemployment experience rather than working experience. The combination of the variables age and unemployment experience, however, (implicitly) controls for the corresponding work experience as well. Finally, human capital theory suggest using the time persons spent with their current employer (firm tenure) as a proxy for firm-specific investment in human capital. Since firm tenure (and its square) is more closely related to labour productivity than the general working experience it should cause an extra increase in wages.

To account for the potential correlation between the kind of job an individual holds and her/his health status seven dummies covering the occupational status are included.[4] In order to further control for other structural factors that may affect wages, I control for sector and time fixed effects as well as other binary variables distinguishing between the eastern and western part of Germany, full-time and part-time employment, and German versus non-German nationality.

Thus, enhancing equation (3.3) according to the previous discussion yields:

$$\log w_{i,t} = \mathbf{b}_{s,t}\boldsymbol{\alpha} + \mathbf{f}_{i,t}\boldsymbol{\beta} + \mathbf{a}_{\mathbf{i},\mathbf{t}}\boldsymbol{\gamma} + \mathbf{ue}_{i,t}\boldsymbol{\upsilon} + \mathbf{ft}_{i,t}\boldsymbol{\tau} + \delta f(h_{i,t}) + \mathbf{du}_{i,t}\boldsymbol{\pi} + error, \quad (3.4)$$

where $\mathbf{b}_{s,t}$, $\mathbf{f}_{i,t}$, $\mathbf{a}_{\mathbf{i},\mathbf{t}}$, s_i, and $f(h_{i,t})$ are defined as above; the vector $\mathbf{ue}_{i,t}$ stands for unemployment experience and its square, $\mathbf{ft}_{i,t}$ is the length of time (and its

[4]Since it is likely that the state of health depends on the kind of job one holds, interaction terms between the occupational status and the health variables were included. However, the interaction terms turned out to be statistically insignificant and were, therefore, excluded from the final model. In another specification, I interacted age and health since it seems obvious that the later changes in the course of life time. However, again I did not find any significant results with respect to the interaction terms.

square) a person spent with her/his current employer, and the $\mathbf{du}_{i,t}$ are sector, occupation, part-time work, nationality, and time dummies.

3.3 Econometric Approach

To simplify the notation in this section, the explanatory variables in (3.4) are approximated by the vector $\mathbf{x}_{i,t}$. The basic framework for the discussion is a linear unobserved regression model of the form:

$$w_{i,t} = \beta_0 + \mathbf{x}_{i,t}\boldsymbol{\beta} + c_i + u_{i,t}, \quad t = 1, 2, ..., T; \quad i = 1, 2, ..., N, \tag{3.5}$$

where $\mathbf{x}_{i,t}$ is $1 \times K$, $\boldsymbol{\beta}$ is the $K \times 1$ parameter vector of interest, c_i contains unobserved individual characteristics (genetic endowment, ability to deal with health problems, talents, etc.), and $u_{i,t}$ is an unobserved error term. Correlation between the individual effect c_i and $\mathbf{x}_{i,t}$ causes the well known omitted variable bias problem. A common way to get rid of this problem is the so called *within* or *fixed effects* estimator. It is the pooled OLS estimator from the regression of the time-demeaned $w_{i,t}$ on the equally transformed $\mathbf{x}_{i,t}$. If a balanced panel is available, and for N relatively large compared to T, the conditional mean independence assumption,

A. 1 $E(u_{i,t} \mid \mathbf{x}_{i,1}, \mathbf{x}_{i,2}, ..., \mathbf{x}_{i,T}, c_i) = 0, \qquad t = 1, 2, ..., T,$

is a sufficient condition for the within-estimator to be consistent as T is constant and $N \to \infty$. Assumption A. 1 also states that the $\mathbf{x}_{i,t}$ are *strictly exogenous* conditional on c_i, which is another way of expressing that the disturbance term $u_{i,t}$ is uncorrelated with the explanatory variables in each time period ($E(\mathbf{x}'_{i,s}u_{i,t}) = \mathbf{0}$, $s \neq t$, and $s, t = 1, 2, ..., T$). Under the standard rank condition that rank $(E(\tilde{\mathbf{X}}_i'\tilde{\mathbf{X}}_i)) = K$ the within estimator is defined as:

$$\hat{\boldsymbol{\beta}}_{within} = (\sum_{i=1}^{N} \tilde{\mathbf{X}}_i'\tilde{\mathbf{X}}_i)^{-1}(\sum_{i=1}^{N} \tilde{\mathbf{X}}_i'\tilde{\mathbf{w}}_i) = (\sum_{i=1}^{N}\sum_{t=1}^{T} \tilde{\mathbf{x}}_{i,t}'\tilde{\mathbf{x}}_{i,t})^{-1}(\sum_{i=1}^{N}\sum_{t=1}^{T} \tilde{\mathbf{x}}_{i,t}'\tilde{w}_{i,t}), \tag{3.6}$$

where $\tilde{\mathbf{X}}_i = \mathbf{J}\mathbf{X}_i$, $\tilde{\mathbf{w}}_i = \mathbf{J}\mathbf{w}_i$ ($\mathbf{w}_i$ is $T\times 1$, $\mathbf{X}_i$ is $T\times K$, and $\mathbf{J} = \mathbf{I}_T - \mathbf{i}_T(\mathbf{i}_T'\mathbf{i}_T)^{-1}\mathbf{i}_T'$), and $\tilde{\mathbf{x}}_{i,t} = \mathbf{x}_{i,t} - T^{-1}\sum_{z=1}^{T}\mathbf{x}_{i,z}$, $\tilde{w}_{i,t} = w_{i,t} - T^{-1}\sum_{z=1}^{T} w_{i,z}$.

3.3.1 Panel attrition under conditional mean independence assumption

If a complete panel is available, estimation of equation (3.6) is straightforward. However, in the GSOEP the number of observations differ over years, i.e. not all relevant variables are observed for each person and each time period under consideration. In the study at hand, two causes for missing observations can be distinguished: 1) individuals are not willing to report information with respect to one of the explanatory variables or the dependent variable (item non-response); 2) individuals endogenously decide to participate in the labour market (self-selection). Under these circumstances the conditional mean independence assumption A. 1 becomes:

A. 2 $E(u_{i,t} \mid \mathbf{x}_i, \mathbf{s}_i, \mathbf{d}_i, c_i) = 0, \qquad t = 1, 2, ..., T,$

where $\mathbf{x}_i = (\mathbf{x}_{i,1}, \mathbf{x}_{i,2}, ..., \mathbf{x}_{i,T})$; $\mathbf{s}_i = (s_{i,1}, s_{i,2}, ..., s_{i,T})$ are selection dummies denoting whether an individual i is participating in the labour market at time t, and $\mathbf{d}_i = (d_{i,1}, d_{i,2}, ..., d_{i,T})$ are binary variables indicating item non-response. A. 2 is valid if the $(\mathbf{s}_i, \mathbf{d}_i)$ are strictly exogenous conditional on c_i and $\mathbf{x}_i$. Assumption A. 2 allows $(\mathbf{s}_i, \mathbf{d}_i)$ to be correlated with c_i or $\mathbf{x}_i$. That is, for the within-estimator to be consistent, it is not necessary that selection into or out of the data set is completely random.

Under the further condition that $\sum_{t=1}^{T} E(s_{i,t}d_{i,t}\tilde{\mathbf{x}}'_{i,t}\tilde{\mathbf{x}}_{i,t})$ is non-singular, pooled OLS on the unbalanced panel yields the following parameter vector:

$$\hat{\boldsymbol{\beta}}_{within} = (\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{x}}'_{i,t}\tilde{\mathbf{x}}_{i,t})^{-1}(\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{x}}'_{i,t}\tilde{w}_{i,t}), \tag{3.7}$$

where $\tilde{\mathbf{x}}_{i,t} = \mathbf{x}_{i,t} - T_i^{-1}\sum_{z=1}^{T} s_{i,z}d_{i,z}\mathbf{x}_{i,z}$, $\tilde{\mathbf{w}}_{i,t} = \mathbf{w}_{i,t} - T_i^{-1}\sum_{z=1}^{T} s_{i,z}d_{i,z}\mathbf{w}_{i,z}$, and $T_i = \sum_{z=1}^{T} s_{i,t}d_{i,t}$.

3.3.2 Selection correction in unobserved effects models

The within estimator of section 3.3.1 is a reasonable approach when we can be sure that condition A. 2 holds. If the decision to participate in the labour market $\mathbf{s}_i$ is, however, correlated with $u_{i,t}$, the estimator in (3.7) is inconsistent. That means, the participation decision is neither randomly determined nor fully covered by some of the observable variables.

In the study at hand, I consider health as an determinant of wages and labour supply, and I am interested in making statements about the impact of health on wages for *all* individuals. Sample selection arises if some unobservable components of the working decision also affect wages. In this respect, one could think of the genetic endowment and the life situation of an individual (e.g. alcohol and nicotine consume, (un)healthy lifestyle, sport activities, etc.). It is a natural assumption that genetic conditions are time-invariant, whereas the personal life situation is likely to change in the course of time. Consequently, for the former, the relationship between the selection process and wages can be completely described by an individual specific fixed effect. The later, on the other hand, is time-variant and for this reason not covered by c_i. As a result, the selection effect of an individual's life situation is influencing wages through the error term $u_{i,t}$. Since these factors are also correlated with health as an explanatory in the wage equation, the failure to control for the selection process may lead to inconsistent estimation.

To overcome the selection problem, the following model is estimated:

$$w_{i,t} = \beta_0 + \mathbf{x}_{i,t}\boldsymbol{\beta} + c_i + u_{i,t}, \quad t = 1, 2, ..., T; \quad i = 1, 2, ..., N, \tag{3.8}$$

$$s^*_{i,t} = \gamma_0 + k_i + \mathbf{z}_{i,t}\boldsymbol{\gamma} + e_{i,t}, \tag{3.9}$$

$$s_{i,t} = \begin{cases} 1 & \text{if } e_{i,t} > -\gamma_0 - \mathbf{z}_{i,t}\boldsymbol{\gamma} - k_i \\ 0 & \text{otherwise,} \end{cases} \tag{3.10}$$

where (3.8) equals (3.5), (3.9) and (3.10) describe a person's decision to participate in the labour market, $s^*_{i,t}$ is the latent propensity to work, $\mathbf{z}_{i,t}$ is a $1 \times G$ vector of covariates, and $\boldsymbol{\gamma}$ is the corresponding parameter vector ($G \times 1$). The variable $w_{i,t}$ is only observed when $s_{i,t} = d_{i,t} = 1$, and the $(\mathbf{z}_{i,t}, s_{i,t})$ are observable for $d_{i,t} = 1$.[5] It is usually assumed that $G > K$, meaning that $\mathbf{z}_{i,t}$ includes at least one exogenous variable that identifies selection. The individual effect k_i contains unobserved characteristics and exhibits no variation over time. Furthermore, $e_{i,t}$, which is normally distributed with standard deviation σ^e_t, is uncorrelated with k_i, $\mathbf{z}_i = (\mathbf{z}_{i,1}, ..., \mathbf{z}_{i,T})$, and $\mathbf{d}_i = (\mathbf{d}_{i,1}, ..., \mathbf{d}_{i,T})$. Following Mundlak (1978), Chamberlain (1984), and Wooldridge (1995) the time-invariant effects are assumed to be linked with $\mathbf{z}_{i,t}$ through a linear function of k_i on the time averages of $\mathbf{z}_{i,t}$ (denoted as $\bar{\mathbf{z}}_i$) and an error term a_i, that is independent of $(\mathbf{z}_i, \mathbf{d}_i)$ and $e_{i,t}$. Equation (3.9) therefore becomes:

$$s^*_{i,t} = \gamma_0 + \psi_0 + \bar{\mathbf{z}}_i\boldsymbol{\psi} + \mathbf{z}_{i,t}\boldsymbol{\gamma} + a_i + e_{i,t} = \theta_0 + \bar{\mathbf{z}}_i\boldsymbol{\theta} + \mathbf{z}_{i,t}\boldsymbol{\gamma} + v_{i,t}, \tag{3.11}$$

[5]In the case of item non-response ($d_{i,t} = 0$), the corresponding observation is missing in both the selection and the main equation.

where $\theta_0 = \gamma_0 + \psi_0$; $\boldsymbol{\theta} = \boldsymbol{\psi}$, $\boldsymbol{\theta}$ and $\boldsymbol{\psi}$ are $G \times 1$ parameter vectors, and a_i is zero mean normally distributed. The distribution of the composite error term $v_{i,t} = a_i + e_{i,t}$ is normal with standard deviation $\sigma_t^v = \sigma^a + \sigma_t^e$. It is allowed to be heterogeneously distributed over time and there are no restrictions imposed on the correlation between $v_{i,t}$ and $v_{i,s}$, i.e. $Cov(v_{i,t}, v_{i,s}) \neq 0$ for $s \neq t$.

Implicitly, assumptions on the selection equations (3.9) and (3.11) were already mentioned in the above, but I summarise them in the following (see also Wooldridge (1995), p.126, and Dustmann and Rochina-Barrachina (2000), p.6):

A. 3 *The unobserved effect in the selection equation can be described as a linear projection of k_i on $\bar{\mathbf{z}}_i$, where $\bar{\mathbf{z}}_i = P_i^{-1} \sum_{s=1}^T d_{i,s}\mathbf{z}_{i,s}$, and $P_i^{-1} = \sum_{s=1}^T d_{i,s}$.*

A. 4 *The errors $v_{i,t} = a_i + e_{i,t}$ are independent of $(\mathbf{z}_i, \mathbf{d}_i)$ and they are normally distributed, $N(0, \sigma_t^v)$.*

The next step is to estimate equation (3.11) using standard probit for each t and obtain the inverse Mills ratios (IMRs) for $s_{i,t} = d_{i,t} = 1$ as $\hat{\lambda_{i,t}} = \phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)/\Phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)$, where $\mathbf{h}_{i,t} = (1, \bar{z}_{i,1}, ..., \bar{z}_{i,G}, z_{i,1,t}, ..., z_{i,G,t})$ and $\hat{\boldsymbol{\delta}}_t = (\hat{\theta_{0,t}}, \hat{\theta_{1,t}}, ..., \hat{\theta_{G,t}}, \hat{\gamma_{1,t}}, ..., \hat{\gamma_{G,t}})'$. At this point, it seems tempting to include the IMRs as additional regressors and to estimate equation (3.8) using the within-estimator described in (3.7). However, as Wooldridge (2002) points out, this is (usually) not a valid strategy to arrive at consistent estimates.[6] Instead, he suggests a method that allows the selection term $\hat{\lambda_{i,t}}$ to be *not* strictly exogenous in (3.8) (i.e there are no restrictions on how $u_{i,t}$ relates to $v_{i,s}$, $s \neq t$).[7]

[6] It is, however, possible to use the Within estimator for testing purposes. Under the null hypothesis in A. 2, the IMRs should not be significant when using the within-estimator on an augmented version of equation (3.4). See also section 3.5.

[7] To place more emphasis on this, without abandoning the strict exogeneity assumption for the IMR at this point it is not possible to allow for serial correlation in the selection equation.

This strategy necessitates to specifically model the unobserved effect such that correlation between c_i and $(\mathbf{x}_i, v_{i,t})$ is possible. Explicitly, the assumptions are:

A. 5 $E(u_{i,t} \mid \mathbf{z}_i, \mathbf{d}_i, v_{i,t}) = E(u_{i,t} \mid v_{i,t}) = L(u_{i,t} \mid v_{i,t}) = \rho_t v_{i,t}$,

i.e. $u_{i,t}$ is mean independent of $(\mathbf{z}_i, \mathbf{d}_i)$ conditional on $v_{i,t}$ and the conditional mean of $u_{i,t}$ is a linear function of $v_{i,t}$.

A. 6 $E(c_i \mid \mathbf{z}_i, \mathbf{d}_i, v_{i,t}) = L(c_i \mid 1, \bar{x}_{i,1}, ..., \bar{x}_{i,K}, v_{i,t}) = \tau_0 + \tau_1 \bar{x}_{i,1} + ... + \tau_k \bar{x}_{i,K} + \varsigma_t v_{i,t}$,

i.e. the unobserved effect in the main equation can be described as a linear projection of c_i on $(\bar{\mathbf{x}}_i,\ v_{i,t})$ and an error term b_i, where $\bar{\mathbf{x}}_i = (\bar{x}_{i,1}, ..., \bar{x}_{i,K})$, and the conditional expectation of b_i is independent of $(\mathbf{z}_i, \mathbf{d}_i)$ and $v_{i,t}$ $(E(b_i \mid \mathbf{z}_i, \mathbf{d}_i, v_{i,t}) = 0)$.

At this point, it seems necessary to spend a few words on the item non-response indicators $\mathbf{d}_i$ in A. 4, A. 5, and A. 6. In the case where item non-response is entirely random or some slightly weaker assumptions (see above), $\mathbf{d}_i$ is independent of $(\mathbf{u}_i,\ \mathbf{s}_i, \mathbf{z}_i, c_i, k_i)$. Hence, $\mathbf{d}_i$ is independent of $\mathbf{v}_i$ in A. 4 and assumptions A. 5 and A. 6 hold under $E(u_{i,t} \mid \mathbf{z}_i, v_{i,t}) = E(u_{i,t} \mid v_{i,t}) = \rho_t v_{i,t}$ and $E(c_i \mid \mathbf{z}_i, v_{i,t}) = L(c_i \mid 1, \bar{\mathbf{x}}_i, v_{i,t}) = \tau_0 + \bar{\mathbf{x}}_i \boldsymbol{\tau} + \varsigma_t v_{i,t}$. However, the assumption of complete randomness is stronger than actually needed. If there is item non-response, the corresponding observation is missing both in the selection and in the main equation. So, one needs to assume that $\mathbf{d}_i$ is independent of the error term $v_{i,t}$ in the participation equation, and conditional mean independent of $u_{i,t}$. Nevertheless, $\mathbf{d}_i$ is still allowed to be correlated with $(\mathbf{z}_i, k_i)$. Since $v_{i,t}$ is a determinant of c_i (see A. 6) $\mathbf{d}_i$ needs to be uncorrelated with the unobserved effect in the main equation.[8]

[8] One should be aware of the fact that the random item non-response assumption might be doubted if persons are not willing or able to reply to the GSOEP due to their poor health status. Unfortunately, it is not possible to control for this eventuality and the random drop-out assumption needs to be maintained at this point.

The conditional expectation for $w_{i,t}$ can, then, be expressed as:

$$\begin{aligned} E(w_{i,t} \mid \mathbf{z}_i, \mathbf{d}_i, v_{i,t}) &= E(w_{i,t} \mid \mathbf{z}_i, v_{i,t}) \\ &= E(c_i \mid \mathbf{z}_i, v_{i,t}) + \beta_0 + \mathbf{x}_{i,t}\boldsymbol{\beta} + E(u_{i,t} \mid \mathbf{z}_i, v_{i,t}) \\ &= (\beta_0 + \tau_0) + \bar{\mathbf{x}}_i\boldsymbol{\tau} + \mathbf{x}_{i,t}\boldsymbol{\beta} + (\varsigma_t + \rho_t)v_{i,t} \\ &= \varphi_0 + \bar{\mathbf{x}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t v_{i,t}. \end{aligned} \tag{3.12}$$

Here, the first and second equality hold under the assumption that item non-response is entirely random (see above); $\varphi_0 = \beta_0 + \tau_0$, $\boldsymbol{\varphi} = \boldsymbol{\tau}$, $\boldsymbol{\varphi}$ and $\boldsymbol{\tau}$ are $K \times 1$ parameter vectors, and $\xi_t = \varsigma_t + \rho_t$.[9] Using the law of iterated expectations on equation (3.12) yields:

$$\begin{aligned} E(w_{i,t} \mid \mathbf{z}_i, s_{i,t}) &= \varphi_0 + \bar{\mathbf{x}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t E(v_{i,t} \mid \mathbf{z}_i, s_{i,t}) \\ &= \varphi_0 + \bar{\mathbf{x}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t f(\mathbf{z}_i, s_{i,t}), \end{aligned} \tag{3.13}$$

where $f(\mathbf{z}_i, s_{i,t})$ is a function of $\mathbf{z}_i$ and $s_{i,t}$. Since in the selected sample $w_{i,t}$ is only observable for $s_{i,t} = 1$, $f(.)$ can be replaced by $f(\mathbf{z}_i, s_{i,t} = 1) = f(\mathbf{z}_i, v_{i,t} > -\mathbf{h}_{i,t}\boldsymbol{\delta}_t) = \phi(\mathbf{h}_{i,t}\boldsymbol{\delta}_t)/\Phi(\mathbf{h}_{i,t}\boldsymbol{\delta}_t) = \lambda_{i,t}$.

As mentioned before, the crucial point is that $v_{i,s}$, for $s \neq t$, is not in the conditioning set of A. 5 and so Wooldridge's estimator allows for serial correlation and heterogeneity in the error terms of the selection equation. Stated

[9] With the exception of the constant term, identifying the vector $\boldsymbol{\beta}$ can easily be achieved since by the law of iterated expectations:

$$\begin{aligned} E(c_i \mid \mathbf{z}_i, \mathbf{d}_i) &= \tau_{0,t} + \bar{\mathbf{x}}_i\boldsymbol{\tau}_t + \rho_t E(v_{i,t} \mid \mathbf{z}_i, \mathbf{d}_i) \\ &= \tau_{0,t} + \bar{\mathbf{x}}_i\boldsymbol{\tau}_t = \tau_0 + \bar{\mathbf{x}}_i\boldsymbol{\tau}. \end{aligned}$$

The second equality holds because $E(v_{i,t} \mid \mathbf{z}_i, \mathbf{d}_i) = 0$ in assumption A. 4 and the third equality follows due to the fact that the coefficients describing the time constant effects are necessarily time-invariant. If variables are not changing over time it is impossible to distinguish β_k and φ_k. Furthermore, there is now way to determine how much of the selection process works through c_i and how much through the time varying unobserved factors in $u_{i,t}$.

differently, $s_{i,s}$, for $s \neq t$, is not in $E(v_{i,t} \mid \mathbf{z}_i, s_{i,t})$ and so the error term $r_{i,t}$ in

$$\begin{aligned} w_{i,t} &= \varphi_0 + \bar{\mathbf{x}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t\lambda_{i,t} + (b_i + l_{i,t}) \\ &= \varphi_0 + \bar{\mathbf{x}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t\lambda_{i,t} + r_{i,t} \end{aligned} \tag{3.14}$$

is allowed to be correlated with $\lambda_{i,s}$, for $s \neq t$, where $l_{i,t}$ is part of the composite error term $u_{i,t} = \rho_t v_{i,t} + l_{i,t}$ and b_i is defined as above. Dustmann and Rochina-Barrachina (2000) call the condition $E(r_{i,t} \mid \mathbf{z}_i, s_{i,t}) = 0$ "contemporaneous exogeneity" of the selection term with respect to $r_{i,t}$.

The simplest way to consistently estimate (3.14) (with $\lambda_{i,t}$ replaced by $\hat{\lambda}_{i,t}$) if strict exogeneity (with respect to the IMRs) fails is pooled OLS. When calculating the asymptotic variance of $(\varphi_0, \boldsymbol{\varphi}', \boldsymbol{\beta}', \boldsymbol{\xi}')'$, I follow Wooldridge (1995) and construct standard errors robust to serial correlation and heteroscedasticity that are also adjusted for the additional variation introduced by the estimation of T probit models in the first step. The calculation of the asymptotic variance covariance estimator is described in appendix C.1.

3.3.3 Panel attrition with endogenous regressors

Estimation of equation (3.14) assumes (strict) exogeneity of the explanatory variables. However, in the study at hand – even after controlling for individual specific heterogeneity and sample selection – the health variable is likely to be endogenous. Three cases of endogenity may appear. 1) Since health satisfaction is a self-assessed variable, measurement error might pose a problem; 2) the health condition may benefit from rising wages as the marginal return of health investment increases with the wage rate (reverse causality); 3) if past shocks affect current health, the health variable is not strictly exogenous in the wage equation.

Semykina and Wooldridge (2005) provide an estimation method based on Wooldridge (1995) that accounts for endogeneity in the presence of unobserved heterogeneity and sample selection. Analogous to section 3.3.1, it seems reasonable to start with a mean independence assumption that allows for consistent estimates in an unbalanced panel framework, when some of the explanatory variables are endogenous. Presume that the health variable (as part of $x_{i,t}$ in equation (3.5)) is correlated with $u_{i,t}$. Furthermore, suppose that a vector of instruments $\mathbf{q}_{i,t}$ $(1 \times E)$ is available, which consists of all exogenous variables in $\mathbf{x}_{i,t}$ and at least one instrument.[10] Then, for the Within- or FE-2SLS (two step least square) estimator in an unbalanced panel framework to be consistent, the equivalent to A. 2 is:

A. 7 $E(u_{i,t} \mid \mathbf{q}_i, \mathbf{s}_i, \mathbf{d}_i, c_i) = 0, \qquad t = 1, 2, ..., T,$

where $\mathbf{q}_i = (\mathbf{q}_{i,1}, \mathbf{q}_{i,2}, ..., \mathbf{q}_{i,T})$, $\mathbf{q}_{i,t} = (q_{i,t,1}, ..., q_{i,t,E})$, and the $(\mathbf{s}_i,\mathbf{d}_i)$ are defined as in section 3.3.1. A. 7 requires sample attrition $(\mathbf{s}_i,\ \mathbf{d}_i)$ and the vector of instruments $\mathbf{q}_i$ to be strictly exogenous conditional on c_i. Moreover, all variables in $\mathbf{q}_i$ are assumed to vary over time, $\mathbf{q}_i$ is allowed to be correlated with c_i, and the $(\mathbf{s}_i,\mathbf{d}_i)$ are either completely random or a function of $(\mathbf{q}_i, c_i)$. If there are no linear dependencies among the demeaned $\mathbf{q}_{i,t}$ (rank $E(\sum_{t=1}^{T} s_{i,t} d_{i,t} \tilde{\mathbf{q}}'_{i,t} \tilde{\mathbf{q}}_{i,t}) = E$, $\tilde{\mathbf{q}}_{i,t} = \mathbf{q}_{i,t} - T_i^{-1} \sum_{z=1}^{T} s_{i,z} d_{i,z} \mathbf{q}_{i,z}$, and $T_i = \sum_{z=1}^{T} s_{i,z} d_{i,z}$) and if rank $E(\sum_{t=1}^{T} s_{i,z} d_{i,t} \tilde{\mathbf{x}}'_{i,t} \tilde{\mathbf{q}}_{i,t}) = K$ (i.e. the instruments are partially correlated with the endogenous variables conditional on

[10] It is assumed that unemployment, experience, and years of schooling in equation (3.4) are strictly exogenous conditional in c_i.

the exogenous part of $\mathbf{x}_{i,t}$) the FE-2SLS estimator is given by:

$$\begin{aligned}\hat{\boldsymbol{\beta}}_{FE-2SLS} &= [(\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{x}}_{i,t}'\tilde{\mathbf{q}}_{i,t})'(\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{q}}_{i,t}'\tilde{\mathbf{q}}_{i,t})^{-1} \\ &\quad (\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{q}}_{i,t}'\tilde{\mathbf{x}}_{i,t})]^{-1} \times (\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{x}}_{i,t}'\tilde{\mathbf{q}}_{i,t})' \\ &\quad (\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{q}}_{i,t}'\tilde{\mathbf{q}}_{i,t})^{-1}(\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\tilde{\mathbf{q}}_{i,t}'\tilde{w}_{i,t}). \qquad (3.15)\end{aligned}$$

As in section 3.3.2, it is assumed that item non-response occurs randomly. That means, condition A. 7 alters to:

A. 8 $E(u_{i,t} \mid \mathbf{q}_i, \mathbf{s}_i, \mathbf{d}_i, c_i) = E(u_{i,t} \mid \mathbf{q}_i, \mathbf{s}_i, c_i) = 0, \qquad t = 1, 2, ..., T.$

3.3.4 Selection correction in unobserved effects models with endogeneity

The final step is to derive an estimator that allows $v_{i,t}$ in (3.11) to be correlated with $u_{i,t}$ and c_i in (3.8), when the health variable is endogenous (meaning that $E(r_{i,t} \mid \mathbf{x}_i, s_{i,t}) \neq 0$ in equation (3.14)).

Consider a model that consists of the main equation (3.8) and a selection process that occurs according to the following equation:

$$s_{i,t}^{*} = \gamma_0 + k_i + \mathbf{q}_{i,t}\boldsymbol{\gamma} + e_{i,t}; \qquad s_{i,t} = 1[s_{i,t}^{*} > 0], \qquad (3.16)$$

where 1[.] is an indicator function that equals one if its argument is true, and zero otherwise. Again, the selection equation rests on assumptions A. 3 and A. 4, except that now the $1 \times G$ vector $\bar{\mathbf{z}}_i$ and the $1 \times TG$ vector $\mathbf{z}_i$ are replaced by the $1 \times E$ vector $\bar{\mathbf{q}}_i$ and the $1 \times TE$ vector $\mathbf{q}_i$. Under these assumptions, equation (3.16) becomes:

$$s^*_{i,t} = \gamma_0 + \psi_0 + \bar{\mathbf{q}}_i\boldsymbol{\psi} + \mathbf{q}_{i,t}\boldsymbol{\gamma} + a_i + e_{i,t} = \theta_0 + \bar{\mathbf{q}}_i\boldsymbol{\theta} + \mathbf{q}_{i,t}\boldsymbol{\gamma} + v_{i,t}. \tag{3.17}$$

Likewise, A. 5 and A. 6 are imposed on the relationship between the selection process and $(u_{i,t}, c_i)$, where the vector $\mathbf{z}_i$ is replaced by $\mathbf{q}_i$ and $\bar{x}_{i,j}$, $j = 1, ..., K$, is now $\bar{q}_{i,p}$, $p = 1, ..., E$. Then, the conditional expectation in (3.12) can be rewritten as:

$$E(w_{i,t} \mid \mathbf{q}_i, v_{i,t}) \;\; = \;\; \varphi_0 + \bar{\mathbf{q}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t v_{i,t}, \tag{3.18}$$

where $\xi_t = (\varsigma_t + \rho_t)$. Using the law of iterated expectations on (3.18) and plugging into (3.8) yields:

$$\begin{aligned} w_{i,t} &= \varphi_0 + \bar{\mathbf{q}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t E(v_{i,t} \mid \mathbf{q}_i, s_{i,t}) + r_{i,t}, \\ &= \varphi_0 + \bar{\mathbf{q}}_i\boldsymbol{\varphi} + \mathbf{x}_{i,t}\boldsymbol{\beta} + \xi_t f(\mathbf{q}_i, s_{i,t}) + r_{i,t}. \end{aligned} \tag{3.19}$$

Again, the first step is to estimate T standard probit models of equation (3.17), and calculate the IMRs $\hat{\lambda}_{i,t}$. Then, because the selected sample has $s_{i,t} = 1$, $f(\mathbf{q}_i, s_{i,t})$ in equation (3.19) can be expressed as $f(\mathbf{q}_i, s_{i,t} = 1) = f_t(\mathbf{q}_i, v_{i,t} > -\mathbf{h}_{i,t}\boldsymbol{\delta}_t) = \phi(\mathbf{h}_{i,t}\boldsymbol{\delta}_t)/\Phi(\mathbf{h}_{i,t}\boldsymbol{\delta}_t) = \lambda_{i,t}$, where $\mathbf{h}_{i,t} = (1, \bar{q}_{i,1}, ..., \bar{q}_{i,E}, q_{i,t,1},$ $..., q_{i,t,E})$ and $\boldsymbol{\delta}_t$ is the corresponding parameter vector. Finally, since $r_{i,t}$ is allowed to be correlated with $\lambda_{i,s}$, for $s \neq t$, (i.e. $\lambda_{i,t}$ is not strictly exogenous in (3.19)), a consistent way of estimating (3.19) – with $f(\mathbf{q}_i, s_{i,t} = 1)$ replaced by $\hat{\lambda}_{i,t}$ – is pooled 2SLS, where $1, \bar{\mathbf{q}}_i, \mathbf{q}_{i,t}, \hat{\lambda}_{i,t}$ serve as instruments ($1, \hat{\lambda}_{i,t}$, and the exogenous variables in $\mathbf{x}_{i,t}$ are used as their "own" instruments).

To calculate the asymptotic variance of $(\hat{\varphi}_0, ..., \hat{\varphi}_E, \hat{\beta}_1, ..., \hat{\beta}_K, \hat{\xi}_1, ..., \hat{\xi}_T)'$, I follow Semykina and Wooldridge (2005) and construct standard errors robust to serial correlation and heteroscedasticity that are adjusted for the additional variation introduced by the estimation of T probit models in the first step. They also account for the use of the pooled 2SLS estimator. The estimation of the asymptotic variance covariance matrix is described in appendix C.1.

At last, it is important to describe, how many instruments are needed in the above procedure. As usual, the vector of instruments consists of all exogenous variables in $\mathbf{x}_{i,t}$ and at least as many instruments as there are endogenous variables. Moreover, for the purpose of clearly identifying the parameter vector in the main equation, at least one additional instrument is required. Thus, in the study at hand a minimum of two instruments should be available.

3.4 Data and Descriptives

The data used in this analysis are made available by the German Socio-Economic Panel Study (GSOEP, see SOEP Group 2001) at the German Institute for Economic Research (DIW), Berlin. It is a representative panel data set of the German population that is drawn on a yearly basis. For the western German states, the GSOEP started with about 12,200 observations in 1984. In June 1990 another 4,400 persons living in the former territory of the German Democratic Republic were interviewed in order to expand the GSOEP to the eastern part of Germany.

For the empirical analysis, observations from all sub-samples, with the exception of sample G ("Oversampling of High Income") between 1995 and 2005, are selected.[11] I extract data on the variables described in the appendix (C.2) and exclude (individual-year) observations from both the selection and the wage equation if there is missing data on any of these variables except wages.

[11] 1995 is chosen as starting point because the variable 'number of doctor visits' is not available in 1994.

The sample is constrained to persons older than 17 and younger than 66 years. I exclude those who are self-employed, self-employed in the agricultural sector, work in family business, are on maternity leave, as well as persons attending military/civilian service, and marginally or irregularly employed persons in any of the years under consideration. Individuals who serve an apprenticeship, trainees, interns, volunteers, aspirants, pensioners, and persons still in education are also removed from the estimation sample.

In this kind of study, it is important to discuss whether to include (severely) handicapped persons in the analysis. Motivated by two arguments, I decided to leave them out of the estimation sample. First, firms might discriminate against handicapped persons, irrespective of their productivity. Therefore, their wages might be artificially low or they might drop out of labour market due to discrimination – something that is not meant to be captured in the selection equation. Second, in Germany severely handicapped persons mainly work at special locations (Behindertenwerkstätten), where they are not paid according to their marginal productivity.

The dependent variable used in the main equation is hourly wages derived from individual gross earnings in the month before the interview divided by 4.3 and information on the actual working time per week. In case actual hours worked fall below contractual hours worked, hourly wages are constructed using the later. Any extra salaries like Christmas or holiday bonuses, 13th monthly pay, or child benefits are not taken into account. When calculating hourly wages suspiciously high or low values were manually overseen and dropped if necessary. Wages (as well as all other financial variables) are deflated to their year 2001 real values using the eastern and western CPIs and – if necessary – converted into euro figures at the rate of 1.95583 (the conversion rate of the

Euro in 1999).[12]

Individuals are defined as participating in the labour market if they work for pay in the month before the interview. In the participation equations both working and non-working persons are used for estimation.[13] After the stepwise exclusion of different groups, I arrive at an estimation sample of $10,081$ female and $9,540$ male persons, resulting in $48,763$ and $48,536$ observations, respectively. For estimating earnings equations, persons who work for only one year are dropped from the sample. Due to this restriction and because observations with missing wages are included in the participation equation, the number of observations in the wage equations (29,304 and 39,048; see tables C.6 and C.5 in the appendix) differs slightly from the working population in the probit sample (30,689 and 40,399; see tables C.4 and C.3).

Tables C.3 and C.4 in the appendix compare variables in the participation equation for working and non-working individuals, and tables C.5 and C.6 depict summary statistics of the variables used in the earnings equations. The health variable, which reports current health satisfaction of individuals, is categorial, ranging from zero to ten. It is transformed using the following log-function: $f(h_{i,t}) = \log(h_{i,t} + \sqrt{(h_{i,t}^2 + 1)})$. Health satisfaction differs between the working and non-working population. On average, the transformed value for working females between 1995 and 2005 is around 2.583, while the value for non-working women is smaller at about 2.49 log points. For males, the working non-working health ratio is about 2.598 to 2.406. The hypothesis of the equality of means between the working and non-working group can be rejected on the basis of two standard t-test, t = 23.38 (p-value = 0) for females and t = 38.73 (p-value = 0) for males. In the time period considered, about 63% of

[12]For this purpose, Consumer Price Indices included in the $pequiv files of the GSOEP were used.

[13]I follow Dustmann and Rochina-Barrachina (2000) and include persons as working if they declare participation, but not wages (and if all explanatory variables are available).

the female sample and around 83% of the male sample population participate in the labour market. Men active in the labour market are on average 2.2 years younger, their school attendance was 1.1 years higher, their non-labour income is lower and they have more children than their non-working counterparts. At the same time, a lower portion of male labour market participants is single (21% vs. 32%), has a foreign nationality (12% vs. 19%), and less male workers live in the eastern part of Germany (24% vs. 37%). For women, in this respect the opposite holds true: A larger part of working females live in eastern Germany (28% vs 21%), a smaller portion is married/has a partner (75% vs. 85%), and they have less children than working females. Just like their male colleagues, female workers are slightly younger (2.75 years), spent more time in education or training (1.16 years), and have a lower non-labour income compared to the sample population of female non-workers. Finally, when comparing women and men it becomes clear that in the sample period male real hourly wages were on average about 0.22 log points higher than those of women.

3.5 Empirical Results

Equation (3.4) is estimated using six different estimation methods and tables 3.2 and 3.3 report the results. Pooled OLS in column (1) assumes that the explanatory variables are uncorrelated with individual heterogeneity. If an individual's genetic endowment affects health positively and if it is at the same time more likely to be in the labour market, then OLS estimates should be upward biased. The Within estimator in column (2) helps to overcome this problem as it allows for correlation between health satisfaction and unobserved heterogeneity. The upward bias should be even larger if (positively correlated) time-variant unobservables, determining wages and participation, cause a sample selection problem. An example in this respect is a person's individual life situation characterised by her/his alcohol and nicotine consume, sport activities, healthy lifestyle, etc.. Consequently, in column (3) Wooldridge's (1995) estimator is presented. It allows controlling for heterogeneity and selection in one common framework. However, as argued before, it is unlikely that the health variable is strictly exogenous in the wage equation. A solution to this problem is to use instrumental variable techniques. The pooled 2SLS estimator in column (4) assumes all exogenous variables and the instruments to be uncorrelated with the unobserved effects, whereas the FE-2SLS in column (5) additionally allows for correlated fixed effects. Finally, Semykina's and Wooldridge's (2005) estimator (column (6)) deals with heterogeneity, selection and endogeneity in one estimation approach.

The set of instruments consists of nine variables which also serve as exclusion restrictions in the participation equations: non-labour income, a binary variable for having a partner/being married, age of the partner/spouse and its square, labour market experience of the partner/spouse and its square, and education (in years) of the partner/spouse and its square.[14] Furthermore, I use two extra instruments that are not included in the selection equation. 1) the number of doctor visits in the last three months; 2) the number of days absent from work due to illness in the last year. At this juncture, the argument is that both variables approximate past investment (and depreciation) in health and account for past shocks affecting current health satisfaction. To check the rank conditions on the 2SLS estimators, F-tests on the joint-significance of the instruments in the first step regressions are conducted. For both women and

[14] Wooldridge (2002) suggests to add all exogenous variables, which appear in the selection equation, to the list of instruments. He argues that it can be 'dangerous' to introduce any exclusion restrictions up on reduced form equations. However, based on the prior information that some authors find a direct relationship between the number of children and wages, I do not use these kind of variables as instruments, though they are excluded from the earnings equations and included in the participation equations.

men and for all econometric models the null hypotheses are rejected at any sensible level.

The presumption that a selection bias exists is testable. In table 3.1 a number of Wald tests on the joint significance of 11 Inverse Mills Ratios, each one constructed using a separate probit, are provided. In columns (1) and (2) I follow Wooldridge (1995) and conduct so called 'variable addition' tests, that were first proposed by Verbeek and Nijman (1992). It is assumed that no further endogeneity problems occur. Under the null hypothesis the Within estimator in section 3.3.1 is valid. In columns (3) and (4) tests in the spirit of Semykina and Wooldridge (2005) are accomplished. The null hypothesis suggests to use the FE-2SLS estimator of section 3.3.3. Both procedures (as well as all further estimates) are done separately for females and males to account for expected gender differences in wage determination.

As it turns out, for both women and men, the inverse Mills ratios are negatively correlated with wages for most years. Since the IMRs are inversely related to the estimated probabilities of being employed, derived from the first step probit equations, the negative coefficients indicate that a higher participation probability is associated with an above average salary. For men, the test procedures provide evidence on a selection bias in both the Within and the FE-2SLS framework. The χ^2 statistics, with 11 degrees of freedom, are 31.64 and 26.96, respectively, which gives p-values of about 0.001 and 0.005. Interestingly, for women a selection correction is not indicated. The χ^2 statistics for females are 12.65 and 12.42, resulting in p-values of 0.317 and 0.312. Thus, for women the null hypothesis of no selection bias can not be rejected, a results also found in Dustmann and Rochina-Barrachina (2000). This suggests that in the female sample the selection process is already accounted for by

Table 3.1: IMR TESTS, WOMEN AND MEN, 1995-2005

	male FE[a)]	female FE[a)]	male FE-2SLS[b)]	female FE-2SLS[b)]
	(1)	(2)	(3)	(4)
IMR 1995	.046 (.030)	.011 (.022)	.051 (.031)*	.019 (.023)
IMR 1996	-.007 (.021)	.011 (.019)	-.005 (.021)	.017 (.019)
IMR 1997	-.014 (.031)	-.002 (.019)	-.011 (.030)	.004 (.020)
IMR 1998	.009 (.024)	-.002 (.023)	.012 (.023)	.003 (.023)
IMR 1999	-.031 (.018)*	-.0003 (.020)	-.027 (.018)	.003 (.020)
IMR 2000	-.049 (.016)***	-.020 (.017)	-.045 (.016)***	-.016 (.017)
IMR 2001	-.058 (.016)***	-.034 (.016)**	-.052 (.017)***	-.031 (.016)*
IMR 2002	-.020 (.019)	-.023 (.017)	-.014 (.019)	-.020 (.017)
IMR 2003	-.038 (.016)**	-.029 (.019)	-.032 (.016)**	-.024 (.019)
IMR 2004	-.047 (.017)***	-.046 (.020)**	-.042 (.017)**	-.041 (.020)**
IMR 2005	-.043 (.018)**	-.052 (.021)**	-.035 (.019)*	-.048 (.021)**
Wald-test, χ^2_{11} =	31.64	12.65	26.96	12.42
p-values	.001	.317	.005	.333
N	39,048	29,304	39,048	29,304

Source: GSOEP 1995-2005, own calculations. Within and FE-2SLS estimation. Robust standard errors are in parenthesis: * significance at ten, ** at five, and *** at one percent. Robust p-values are reported under the test statistics. *a)* Wald tests on the joint significance of the IMRs are provided. It is assumed that there are no further endogeneity problems. Under the null hypothesis the Within estimator in section 3.3.1 is valid. *b)* Wald test on the joint significance of the IMRs are provided. Under the null hypothesis the FE-2SLS estimator in section 3.3.3 is valid.

the observable variables and the latent effect c_i. No further evidence is found that any unobservable characteristics in the participation equation affect wages through the error term of the main equation.

For males (table 3.2), the parameter of the health variable using pooled OLS (0.043) is higher than the coefficient in the fixed effects model (0.012). Wooldridge's (1995) estimator, in turn, exhibits the lowest coefficient (0.009) under the assumption of no further endogeneity. All estimates are significant at the 1%-5% confidence level. These results suggest that using the FE estimator already accounts for most of the upward bias introduced by the correlation between the health variable and unobserved individual heterogeneity. Control-

Table 3.2: WAGE EQUATION, MEN, 1995-2005

	OLS$^{a)}$	Within$^{a)}$	Wooldr95$^{b)}$	2SLS$^{a)}$	FE-2SLS$^{a)}$	Wooldr05$^{c)}$
	(1)	(2)	(3)	(4)	(5)	(6)
Health sat.	0.043 (0.004)***	0.012 (0.004)***	0.009 (0.004)**	0.088 (0.012)***	0.071 (0.017)***	0.013 (0.023)
Age	0.097 (0.006)***	.	.	0.097 (0.006)***	.	.
Age square	-.002 (0.0001)***	-.002 (0.0002)***	0.0005 (0.00005)***	-.002 (0.0001)***	-.002 (0.0002)***	0.0005 (0.00006)***
Age triple	1.00e-05 (1.18e-06)***	1.00e-05 (1.62e-06)***	-5.83e-06 (5.77e-07)***	1.00e-05 (1.11e-06)***	1.00e-05 (1.63e-06)***	-5.82e-06 (7.00e-07)***
Unempl. exp.	-.048 (0.003)***	-.097 (0.011)***	-.074 (0.014)***	-.047 (0.003)***	-.098 (0.011)***	-.075 (0.014)***
Unempl. exp. sq.	0.003 (0.0003)***	0.004 (0.002)***	0.003 (0.002)*	0.003 (0.0003)***	0.005 (0.002)***	0.003 (0.002)*
Firm tenure	0.013 (0.0005)***	0.004 (0.0007)***	0.007 (0.0009)***	0.013 (0.0005)***	0.005 (0.0007)***	0.007 (0.0009)***
Firm tenure sq.	-.0002 (1.00e-05)***	-.0001 (0.00002)***	-.0002 (0.00002)***	-.0002 (1.00e-05)***	-.0001 (0.00002)***	-.0002 (0.00003)***
Education	0.032 (0.0008)***	.	.	0.032 (0.0007)***	.	.
Dummy Educ.	-.020 (0.004)***	.	.	-.019 (0.004)***	.	.
Part Time	-.103 (0.015)***	-.042 (0.017)**	-.035 (0.021)*	-.100 (0.011)***	-.041 (0.017)**	-.036 (0.021)*
Foreigner	0.01 (0.005)**	.	.	0.009 (0.005)*	.	.
Lg. unempl. (fed. st.)	-.046 (0.004)***	0.021 (0.013)	0.029 (0.018)	-.045 (0.004)***	0.022 (0.013)*	0.03 (0.02)
Lg. vac. (fed. st.)	0.058 (0.004)***	0.01 (0.007)	-.001 (0.01)	0.057 (0.004)***	0.009 (0.007)	-.002 (0.011)
Firm size (<20 employees)$^{d)}$						
20 - 199	0.082 (0.005)***	0.046 (0.006)***	0.034 (0.007)***	0.081 (0.004)***	0.046 (0.006)***	0.034 (0.007)***
200 - 1999	0.147 (0.005)***	0.058 (0.007)***	0.044 (0.009)***	0.146 (0.005)***	0.058 (0.007)***	0.045 (0.009)***
$\geq$ 2000 employees	0.191 (0.005)***	0.067 (0.008)***	0.051 (0.01)***	0.191 (0.005)***	0.067 (0.008)***	0.051 (0.01)***
Firm size missing	0.085 (0.018)***	0.022 (0.016)	0.024 (0.019)	0.083 (0.015)***	0.021 (0.016)	0.026 (0.019)
Region, where person works (Western Germany)						
East Germany	-.262 (0.006)***	-.032 (0.01)***	-.242 (0.011)***	-.262 (0.005)***	-.033 (0.01)***	-.238 (0.011)***
constant	0.361 (0.089)***	.	.	0.235 (0.089)***	.	.
N	39,048	39,048	39,048	39,048	39,048	39,048
d.f.	39,003	32,035	38,980	39,003	32,035	38,970
Wald tests on the joint significance of						
11 IMRs$^{e)}$	.	.	29.62***	.	.	22.04**
10 time dummies	308.26***	265.76***	42.80***	325.46***	262.63***	40.82***
6 occup. dummies	2802.72***	13.26**	750.72***	3063.53***	13,50**	703.23***
9 sector dummies	1301.85***	64.62***	380.13***	1310.33***	64,86***	381.91***
unobs. effects $^{f)}$	.	.	719.17***	.	.	437,86***

Source: GSOEP 1995-2005, own calculations. Standard errors in parenthesis: * significance at ten, ** at five, and *** at one percent. Year, sector, and occupation dummies are included but not reported. *a)* Robust standard errors are provided using the Huber/White/sandwich estimator; *b)* standard errors are robust to serial correlation and heteroscedasticity. They are also adjusted for the first-stage estimation; *c)* robust standard errors as in *b)*, but the 2SLS estimator is used and accounted for; *d)* for dummy variables, the basis categories are given in parenthesis; *e)* a Wald test on the joint significance of the IMRs is conducted; *f)* the χ^2 test statistics for joint significance of $\bar{\mathbf{x}}_i$ or $\bar{\mathbf{q}}_i$ are reported.

Table 3.3: WAGE EQUATION, WOMEN, 1995-2005

	OLS[a)]	Within[a)]	Wooldr95[b)]	2SLS[a)]	FE2SLS[a)]	Wooldr05[c)]
	(1)	(2)	(3)	(4)	(5)	(6)
Health sat.	0.014 (0.005)***	0.005 (0.005)	0.002 (0.005)	0.018 (0.013)	0.047 (0.018)***	0.021 (0.024)
Age	0.071 (0.007)***	.	.	0.071 (0.007)***	.	.
Age sq.	-.001 (0.0002)***	-.001 (0.0003)***	0.0008 (0.00006)***	-.001 (0.0002)***	-.001 (0.0002)***	0.0008 (0.00007)***
Age tr.	7.70e-06 (1.50e-06)***	5.06e-06 (2.02e-06)**	-9.06e-06 (6.72e-07)***	7.69e-06 (1.43e-06)***	5.19e-06 (2.01e-06)**	-9.02e-06 (8.77e-07)***
Unempl. exp.	-.034 (0.003)***	-.116 (0.015)***	-.100 (0.018)***	-.033 (0.003)***	-.116 (0.015)***	-.101 (0.018)***
Unempl. exp. sq.	0.002 (0.0002)***	0.008 (0.002)***	0.007 (0.003)**	0.002 (0.0003)***	0.008 (0.002)***	0.007 (0.003)**
Firm tenure	0.015 (0.0007)***	0.002 (0.001)**	0.005 (0.001)***	0.015 (0.0007)***	0.002 (0.001)**	0.005 (0.001)***
Firm tenure sq.	-.0002 (0.00002)***	-.00005 (0.00003)	-.0001 (0.00004)***	-.0002 (0.00002)***	-.00005 (0.00003)*	-.0001 (0.00004)***
Education	0.039 (0.0009)***	.	.	0.039 (0.0009)***	.	.
du. educ.	-.028 (0.005)***	.	.	-.028 (0.005)***	.	.
Part time	-.048 (0.004)***	-.004 (0.007)	0.004 (0.008)	-.048 (0.004)***	-.004 (0.007)	0.002 (0.008)
Foreigner	0.006 (0.006)	.	.	0.006 (0.007)	.	.
Lg. unempl. (fed. st.)	-.033 (0.005)***	0.006 (0.015)	0.015 (0.02)	-.033 (0.005)***	0.007 (0.015)	0.014 (0.023)
Lg. vac. (fed. st.)	0.026 (0.005)***	-.003 (0.008)	-.019 (0.011)*	0.026 (0.005)***	-.003 (0.008)	-.018 (0.013)
Firm size (<20 employees)[d)]						
20 - 199	0.087 (0.005)***	0.04 (0.007)***	0.036 (0.009)***	0.087 (0.005)***	0.04 (0.007)***	0.036 (0.009)***
200 - 1999	0.133 (0.006)***	0.06 (0.009)***	0.052 (0.011)***	0.133 (0.005)***	0.06 (0.009)***	0.052 (0.011)***
$\geq$ 2000 employees	0.17 (0.006)***	0.061 (0.009)***	0.049 (0.011)***	0.17 (0.006)***	0.061 (0.009)***	0.049 (0.011)***
firm size missing	0.128 (0.02)***	0.057 (0.017)***	0.094 (0.021)***	0.128 (0.017)***	0.056 (0.017)***	0.093 (0.021)***
Region, where person works (Western Germany)						
East Germany	-.224 (0.006)***	-.035 (0.013)***	-.216 (0.012)***	-.224 (0.006)***	-.033 (0.013)***	-.213 (0.012)***
constant	0.707 (0.104)***			0.697 (0.108)***	.	
N	29,304	29,304	29,304	29,304	29,304	29,304
d.f.	29,259	23,544	29,236	29,259	23,544	29,226
Wald tests on the joint significance of						
11 IMRs[e)]	.	.	9.00	.	.	12.58
10 time dummies	79.49***	102.12***	17.97*	82.79***	103.78***	18.22**
6 occup. dummies	2259.16***	29.57***	674.32***	2270.60***	28.87***	633.81***
9 sector dummies	563.88***	30.05***	156.18***	593.48***	30.06***	158.48***
unobs. effects [f)]	.	.	811.42***	.	.	769.38***

Source: GSOEP 1995-2005, own calculations. Standard errors in parenthesis: * significance at ten, ** at five, and *** at one percent. Year, sector, and occupation dummies are included but not reported. *a)* Robust standard errors are provided using the Huber/White/sandwich estimator; *b)* standard errors are robust to serial correlation and heteroscedasticity. They are also adjusted for the first-stage estimation; *c)* robust standard errors as in *b)*, but the 2SLS estimator is used and accounted for; *d)* for dummy variables, the basis categories are given in parenthesis; *e)* a Wald test on the joint significance of the IMRs is conducted; *f)* the χ^2 test statistics for joint significance of $\bar{\mathbf{x}}_i$ or $\bar{\mathbf{q}}_i$ are reported.

ling for selection reduces the coefficient even further, but differences between the FE and the Wooldridge (1995) estimator are small. Turning to the 2SLS models, a comparison of the parameters shows that the coefficients of health satisfaction in columns (1), (2), and (3) are smaller than those in columns (4), (5), and (6), which is expected if there exists a measurement error problem. Yet, within this framework, the parameter ranking follows the same pattern as in the specifications without instruments. The pooled 2SLS parameter exhibits the highest (significant) parameter (0.088). Using the FE-2SLS estimator reduces the coefficient to a value of 0.071. Though insignificantly different from zero, controlling for selection scales the coefficient even further down to 0.013. For the estimators in columns (3) and (6) a Wald test on the joint significance of the $\hat{\boldsymbol{\varphi}}$ was accomplished. In both cases the resulting values of the test statistics are larger than the critical value, indicating correlated individual effects. Selection tests, where now the assumptions under the null hypothesis are more restrictive than those underlying the tests in table 3.1, exhibit χ^2 statistics of 29.62 and 22.04. Thus, the null hypothesis of no selection can again be rejected.

For women too (table 3.3), six different econometric models are presented, but results are less intuitive than in the case of the male sample. As mentioned before, selection corrections are not indicated; Wald tests on the joint significance of the $(\xi_1, ..., \xi_T)$ for the models in columns (3) and (6) confirm this finding. In the specifications without instrumental variables, only pooled OLS brings about a significant result. When considering the different 2SLS estimators, only the fixed effects approach provides a coefficient which is significantly different from zero. The fact that the coefficients in columns (4), (5), and (6) are all larger than those in (1), (2) and (3) may again indicate measurement error problems in the self assessed health variable. The (significant) parameter of health satisfaction in column (5) exhibits a value of 0.047, whereas the pooled OLS coefficient in column (1) lies at 0.014.

Interpreting the results is straightforward: Since both the dependent and the health variable are given in logs, interrelations between the two can be approximated employing elasticities.[15] For males, raising health satisfaction by 10% increases (hourly) wages approximately by 0.09 to 0.88 percent. In the case of females, the increase of the wage rate ranges from about 0.14 to 0.47 percent.

Turning to the other factors affecting earnings, concave wage profiles are found with respect to the time a person spent at the same firm in all specifications and for women and men. Starting, for example, at a value of two years on the job experience, an additional year at the same firm increases female (male) wages by 0.45% (0.6%), when controlling for selection. Given the high unemployment rates in Germany, it is interesting to see that in all models past unemployment periods significantly decrease wages (at an increasing rate). If the coefficient of education is identified, the returns to an additional year of schooling are almost 4% for women and approximately 3.2% for men.

Results for most of the other variables are as expected. For both women and men wages increase at an decreasing rate with age, and working in the eastern part of Germany or being in part-time employment reduces salaries. In the pooled specifications in columns (1) and (4) a larger average number of job seekers per federal sate negatively influences wages, whereas an increasing amount of notified vacancies raises the wage rate. Finally, as for the structural factors effecting wages, I find industry and occupational wage differentials.[16]

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[15] It is implicitly assumed that health satisfaction is a continuous variable. Assessing health as an categorial variable, a 10% rise in health satisfaction equals roughly the increase by one category.

[16] In all models and both for females and males Wald tests confirm the joint significance

Women and men working in large firms ($\geq$ 2000), ceteris paribus, earn significantly more than in medium-sized firms, which in turn earn more than males and females employed in small firms. These effects are still observed when controlling for individual heterogeneity and selection effects, however, the magnitude of the parameters declines.

3.6 Conclusions

In the final chapter of my dissertation, I employ recently developed estimation methods, which control for selection, individual heterogeneity, and endogeneity in one common framework, and apply them to the question whether health has an effect upon wages. There are a number of important links that connect the state of health and earnings. First, health as part of one's human capital affects labour market productivity and hence wages. Second if the rewards to health investment increase in the wage rate health should rise with wages, implying that there exists the problem of reverse causality. Furthermore, as self-reported health satisfaction is used for estimation, it is not possible to assess one's actual health status accurately and measurement error could be a source of bias. Another shortcoming may arise due to the fact that labour market participation is endogenous, where one reason for the endogeneity is an individual's health status. If panel attrition is not a random phenomena but driven by the individual participation decision employing standard methods may result in inconsistent estimation. Finally, since it is likely that unobserved effects (e.g. genetic endowment) are correlated with health the use of panel data techniques is necessary in order to control for a potential omitted variable bias.

In this study reduced form wage equations for women and men augmented

of six occupational and nine sector dummies at any sensible level.

by a variable measuring health satisfaction are estimated. In an attempt to control for unobserved heterogeneity, sample selection, and endogeneity the estimators proposed by Wooldridge (1995) and Semykina and Wooldridge (2005) are applied. Due to the panel structure of the data it is possible to control for unobserved effects. A number of tests provide evidence that for the male sample selection corrections are indicated, while this issue does not cause any problems in the female population. The results show that good health raises wages. For females an increase in health satisfaction by 10% enhances (hourly) wages approximately by 0.14 to 0.47 percent. In the male sample the increase of the wage rate ranges from about 0.09 to 0.88 percent. The health variable is found to suffer from measurement error. For men, applying pooled OLS or pooled 2SLS is accompanied by an upward bias in the health coefficient.

The estimated effects of health on wages work only for contemporaneous changes in health and wages. It is, however, likely that the state of health follows a persistent stochastic process, where the first source of persistence can easily be controlled for by including fixed effects. Non-inclusion of lagged health variables, to account for state dependency as the second reason of persistency, leaves a source of endogeneity in the model, and I try to compensate for it by utilising instrumental variables. Yet, it seems to be a task for the future to estimate a 'complete' model that allows for identifying all potential sources of endoegeneity plus the dynamics of health in one common framework.

Appendix C

C.1 Asymptotic variance-covariance matrices for the estimators in section 3.3[1]

Given the estimated parameter vector $\hat{\boldsymbol{\varrho}}^{OLS} = (\hat{\varphi}_0, \hat{\boldsymbol{\varphi}}', \hat{\boldsymbol{\beta}}', \hat{\boldsymbol{\xi}}')'$ in section 3.3.2, the asymptotic variance is $\text{Avar}(\hat{\boldsymbol{\varrho}}^{OLS}) = \hat{\mathbf{A}}^{-1}\hat{\mathbf{B}}\hat{\mathbf{A}}^{-1}$. Consistent estimators of $\hat{\mathbf{A}}$ and $\hat{\mathbf{B}}$ are:

$$\hat{\mathbf{A}} = N^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\hat{\mathbf{m}}'_{i,t}\hat{\mathbf{m}}_{i,t}, \tag{C.1}$$

where $\hat{\mathbf{m}}_{i,t}$ is $(1, \overline{\mathbf{x}}_i, \mathbf{x}_{i,t}, 0, ..., 0, \hat{\lambda}_{i,t}, 0, ..., 0)$, a $1 \times (1 + 2K + T)$ vector; and

$$\hat{\mathbf{B}} = N^{-1}\sum_{i=1}^{N}\hat{\mathbf{p}}_i\hat{\mathbf{p}}'_i. \tag{C.2}$$

The $(1 + 2K + T) \times 1$ vector $\hat{\mathbf{p}}_i$ is defined as

$$\hat{\mathbf{p}}_i = \hat{\mathbf{j}}_i - \hat{\mathbf{D}}\hat{\mathbf{k}}_i, \quad i = 1, ..., N, \tag{C.3}$$

and $\hat{\mathbf{j}}_i = \sum_{t=1}^{T} s_{i,t}d_{i,t}\hat{\mathbf{m}}'_{i,t}\hat{r}_{i,t}$, where $\hat{r}_{i,t}$ is the OLS residual from equation (3.14). Next, construct the $(1 + 2G)T \times 1$ vector $\hat{\mathbf{k}}_i$ as $(k'_{i,1}, ..., k'_{i,T})'$ and obtain each $k_{i,t}$ by multiplying the estimated information matrix, $I_t(\hat{\boldsymbol{\delta}}_t)$, for each t with the score, $sc_{i,t}(\hat{\boldsymbol{\delta}}_t)$, of the log-likelihood function for person i at time t.[2]The formulas are given in Maddala (1983) or Semykina and Wooldridge (2005) and need to be calculated using $\mathbf{h}_{i,t}$ and $\hat{\boldsymbol{\delta}}_t$, defined in section 3.3.2. Using e.g. the statistical software Stata© allows for a straightforward derivation of the two terms. First, extract the variance-covariance matrix for the T probit models, calculate the inverse and divide it by the number of observations in each participation equation. Second, use the score option for each probit and multiply it with the corresponding $(1+2G) \times 1$ covariate-vector. Third, multiply the two to obtain T $k_{i,t}$ vectors and stack them as described above.

[1]The derivations in this section are based on Wooldridge (1995) and Semykina and Wooldridge (2005).

[2]$(1 + 2G)$ is the number of covariates in each participation equation, see section 3.3.2.

Finally, a consistent estimator for $\hat{D}$ is

$$\hat{\mathbf{D}} = N^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\hat{\mathbf{m}}'_{i,t}\hat{\boldsymbol{\varrho}}'^{OLS}\hat{\mathbf{F}}_{i,t}. \tag{C.4}$$

Here, $\hat{\mathbf{F}}_{i,t}$ is the $(1+2K+T)\times T(1+2G)$ matrix

$$\hat{\mathbf{F}}_{i,t} = \begin{pmatrix} \mathbf{0} & \mathbf{0} & \dots & \mathbf{0} & \mathbf{0} & \dots & \mathbf{0} \\ \mathbf{0} & \mathbf{0} & \dots & \hat{\mathbf{Z}}_{i,t} & \mathbf{0} & \dots & \mathbf{0} \end{pmatrix}, \tag{C.5}$$

where each zero in the first row block is a $(1+2K)\times(1+2G)$ matrix and each zero in the second row block is a $T\times(1+2G)$ matrix. At last, the $T\times(1+2G)$ matrix $\hat{\mathbf{Z}}_{i,t}$, which is in the tth column block of $\hat{\mathbf{F}}_{i,t}$, is defined as $\hat{\mathbf{Z}}_{i,t} = (\mathbf{0}', \mathbf{0}', ..., (\hat{v}_{i,t}\mathbf{h}_{i,t})', \mathbf{0}', ..., \mathbf{0}')'$, where each zero is a $1\times(1+2G)$ vector, and

$$\hat{v}_{i,t} = -\frac{\phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)[\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t\Phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)+\phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)]}{\Phi(\mathbf{h}_{i,t}\hat{\boldsymbol{\delta}}_t)^2}. \tag{C.6}$$

To calculate the asymptotic variance of the coefficient vector $\hat{\boldsymbol{\varrho}}^{2SLS} = (\hat{\varphi}_0, ..., \hat{\varphi}_E, \hat{\beta}_1, ..., \hat{\beta}_K, \hat{\xi}_1, ..., \hat{\xi}_T)'$ in section 3.3.4, define

$$\text{Avar}(\hat{\boldsymbol{\varrho}}^{2SLS}) = N^{-1}(\hat{\mathbf{C}}'\hat{\mathbf{O}}^{-1}\hat{\mathbf{C}})^{-1}\hat{\mathbf{C}}'\hat{\mathbf{O}}^{-1}\hat{\mathbf{B}}\hat{\mathbf{O}}^{-1}\hat{\mathbf{C}}(\hat{\mathbf{C}}'\hat{\mathbf{O}}^{-1}\hat{\mathbf{C}})^{-1}. \tag{C.7}$$

First, use the $1\times(1+E+K+T)$ vector of regressors $\hat{\mathbf{y}}_{i,t} = (1, \overline{\mathbf{q}}_i, \mathbf{x}_{i,t}, 0, ..., 0, \hat{\lambda}_{i,t}, 0, ..., 0)$ and the $1\times(1+2E+T)$ vector of instruments $\hat{\mathbf{n}}_{i,t} = (1, \overline{\mathbf{q}}_i, \mathbf{q}_{i,t}, 0, ..., 0, \hat{\lambda}_{i,t}, 0, ..., 0)$ to calculate

$$\hat{\mathbf{C}} = N^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\hat{\mathbf{n}}'_{i,t}\hat{\mathbf{y}}_{i,t} \quad \text{and} \quad \hat{\mathbf{O}} = N^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T} s_{i,t}d_{i,t}\hat{\mathbf{n}}'_{i,t}\hat{\mathbf{n}}_{i,t}. \tag{C.8}$$

The formula for $\hat{\mathbf{B}}$ is given in (C.2), but its dimension is now $(1+2E+T)\times(1+2E+T)$, and the $(1+2E+T)\times 1$ vector $\hat{\mathbf{p}}_i$ has the form

$$\hat{\mathbf{p}}_i = \sum_{t=1}^{T}(s_{i,t}d_{i,t}\hat{\mathbf{n}}'_{i,t}\hat{r}_{i,t} - \hat{\mathbf{M}}\hat{\mathbf{k}}_i), \tag{C.9}$$

where $\hat{r}_{i,t}$ is the 2SLS residual from equation (3.19).[3] The $(1+2G)T \times 1$ vector $\hat{\mathbf{k}}_i$ is constructed as described above. $\hat{\mathbf{M}}$, a $(1+2E+T) \times (1+2G)T$ matrix, has the form $\hat{\mathbf{M}} = N^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} s_{i,t} d_{i,t} \hat{\mathbf{n}}'_{i,t} \hat{\boldsymbol{\varrho}}'^{2SLS} \nabla_\delta \hat{\mathbf{y}}'_{i,t}$. Finally, define $\nabla_\delta \hat{\mathbf{y}}'_{i,t}$ like $\hat{\mathbf{F}}_{i,t}$ in (C.5), except that now each of the T zeros in the first row block is $(1+K+E) \times (1+2G)$ and each zero in the second row block is $T \times (1+2G)$. The $T \times (1+2G)$ matrix $\hat{\mathbf{Z}}_{i,t}$, which is in the tth column block of $\nabla_\delta \hat{\mathbf{y}}'_{i,t}$, has the form

$$\hat{\mathbf{Z}}_{i,t} = \begin{pmatrix} -\mathbf{h}_{i,t} \hat{\lambda}_{i,t} (\mathbf{h}_{i,t} \hat{\boldsymbol{\delta}}_t + \hat{\lambda}_{i,t}) \\ \mathbf{0} \\ \cdots \\ \mathbf{0} \end{pmatrix}. \tag{C.10}$$

C.2 Description of variables

Variable	Description
Probit	dummy variable indicating participation in the labour market (probit = 1) or no participation (probit = 0)
Log hourly wage	log gross hourly real wage (deflated to 2001 Euros)
Health satisfaction	variable indicating current health satisfaction of an individual; categories range from $0-10$; transformation: $f(h_{i,t}) = \log(h_{i,t} + \sqrt{(h_{i,t}^2+1)})$
Age	age in years
Unemployment experience	length of unemployment in a person's career; in years, with months in decimal form
Firm tenure	length of time with firm; in years, with months in decimal form
Education	amount of education or training in years

(continued)

[3] Note that $r_{i,t}$ is not the residual from the second stage OLS regression. Instead, it is defined as $\hat{r}_{i,t} = w_{i,t} - \hat{\mathbf{y}}_{i,t} \hat{\boldsymbol{\varrho}}^{2SLS}$.

Variable	Description
Dummy education	after intensive checks, wrong values of the education variable are changed to their maximum (du. educ. = 1)
Part-time	dummy variable indicating part-time work
Foreigner	dummy variable indicating non-German nationality
Log unemployment[a]	(log) yearly averages of job seekers (per federal state)
Log vacancies	(log) notified vacancies (per federal state)
Firm size	four dummy variables indicating different firm sizes; categories: up to 20 employees ; $20 - 199$ employees; $200 - 1999$ employees; larger than 2000 employees
Region live/work	dummy variables indicating where a person lives (probit equ.) or works (wage equ.); Region = 0 if Western Germany
Occupation	seven occupation dummies, constructed using the Erikson, Goldthorpe Class Category IS88 (basis: high serv.)
Sector	ten aggregated sector dummies, based on the NACE classification (basis: agric., forestry, fishing)
Time	eleven time dummies (1995 - 2005) (basis: 1995)
Number of children	no. of children in three categories; 1) up to 2 years old; 2) between 3 - 5 years old; 3) between 6 - 16 years old
Non labour income	household income minus net wage income (in 2001 Euros)
No. of visits doctor	number of doctor visits last three months
No. days off	number of days absent from work due to illness last year; the variable is set to zero if a persons was not working last year
Partner or Spouse variables	
Single	dummy variable indicating whether a person has a partner/is married (single = 0)
Net wage[b]	net wage of partner or spouse
Age	age in years of partner or spouse
Experience	labour market experience of partner/spouse
Education	amount of education or training in years of partner/spouse

[a]Both unemployment and vacancy figures are extracted from Arbeitsstatistik 2005 - Jahreszahlen, provided by the Federal Employment Agency, Nuremberg.

[b]All partner/spouse variables equal zero, if single = 1.

C.3 Participation equations

Tables C.1 and C.2 present estimation results for the participation equations (see equations (3.9), (3.10), and (3.11) in section 3.3.2) between 1995 and 2005 using pooled and 'traditional' random effects probit models and two Mundlak (1978) versions of Chamberlain's (1980) random effects probit model. Columns (3) and (4) depict results where the unobserved effects, k_i, are written as linear predictions on the means of all regressors and an error term a_i, which is assumed to be independent of $\mathbf{h}_{i,t}$ with (constant) variance σ_a^2. This explicitly allows some of the regressors to be correlated with the individual effects (k_i), but means that coefficients of time-invariant regressors, like education, are not identified. Under the further assumption that the participation indicators $(s_{i,1}, ..., s_{i,T})$ are independent conditional on $(\mathbf{h}_i, a_i)$, a random effects probit model is estimated; results are depicted in column (4).[4] The pooled probit model in column (3) (where again the unobserved effects are parameterised using the (within) means of the regressors) offers an estimation approach under less restrictive assumptions. Here, the independence assumption with respect to $(s_{i,1}, ..., s_{i,T})$ can be relaxed. However, a robust variance covariance matrix estimator is required to account for the fact that observations are correlated within individuals over time.[5] Equivalently, in columns (1) – pooled probit – and (2) – random effects probit – the same specifications are considered, but here it is assumed that the unobserved effects, k_i, are uncorrelated with any of the regressors.

The estimated coefficients of the health variable show that for both women and men good health significantly increases the probability to work. To compare the different results, I calculate probability differences of being in (very) good health versus suffering from poor health. On this account, participation probabilities of 'average' individuals are predicted, where persons differ only with respect to their state of health. For a healthy man, using pooled probit (column (1)), the probability to work is $P(s = 1|health = 10, \overline{\mathbf{h}}) - P(s = 1|health = 0, \overline{\mathbf{h}}) = 43$ percentage points higher than for a unhealthy male person. Estimating the same model, but

[4]For a detailed description of the different estimators and corresponding assumptions see Wooldridge (2002), chapter 15.8.

[5]Wald tests for the joint significance of the $\boldsymbol{\theta}$ coefficients confirm the presence of correlated unobserved effects. The resulting values of the test statistic in columns (3) and (4) are for both women and man larger than the critical value of the χ^2 at the one percent level.

controlling for correlated individual effects (column (3)), strongly reduces the probability difference to 8.5 percentage points. In the random effects specification of column (2) the probability to work is 3.6 percentage points higher for healthy than for unhealthy men. Here, controlling for correlated fixed effects results in a probability difference of a single percentage point (column (4)). For women the impact of health satisfaction on labour market participation is also positive and significant in all econometric models. A comparison of healthy and unhealthy females results in probability differences of about 36 percentage points, when the pooled probit estimator without correlated individual effects is considered, and 5.6 percentage points, when controlling for the interaction between individual effects and the health variable. In the random effects models the corresponding values (columns (2) and (4)) are around 9.5 and 3.6 percentage points, respectively.

Results for most of the other variables are as expected. For both women and men, the participation probability increases with age (at an decreasing rate) and education. Living in the eastern part of Germany, being of non-German origin, and the amount of non labour income has a negative effect on the probability to work. Interestingly, many of the partner and children variables exhibit the same sign for women and men. For both sexes, the number of children in different age categories mostly reduce the participation probability. The partner's net wage and her/his labour market experience is associated with a decreasing working probability in most specifications for both females and males.

Table C.1: PARTICIPATION EQUATION, MEN, 1995-2005

	Pooled[a)]	Random Effects[b)]	Mundlak, pooled[a)c)]	Mundlak, R.E.[b)c)]
	(1)	(2)	(3)	(4)
Age	.117 (.041)***	.013 (.059)	.	.
Age square	-.002 (.001)**	.002 (.001)	-.0001 (.0002)	.0008 (.0004)**
Age triple	1.00e-05 (8.45e-06)	-.00004 (1.00e-05)***	-7.05e-06 (2.73e-06)***	-.00003 (4.09e-06)***
Education	.102 (.007)***	.211 (.012)***	. .	. .
Dummy Education	-.038 (.032)	-.095 (.041)**	.	.
Foreigner	-.221 (.045)***	-.431 (.072)***	.	.
Health Sat.	.472 (.024)***	.445 (.032)***	.123 (.022)***	.222 (.036)***
Non labour inc.	-.0004 (.00002)***	-.0009 (.00002)***	-.0005 (.00003)***	-.001 (.00002)***
Number of children				
up to 2 years old	-.026 (.037)	-.072 (.056)	-.119 (.037)***	-.158 (.062)**
between 3 - 5	-.036 (.034)	-.027 (.049)	-.078 (.034)**	-.084 (.055)
between 6 - 16	-.060 (.020)***	-.094 (.028)***	-.081 (.024)***	-.112 (.035)***
Partner/Spouse variables				
Single	3.502 (.403)***	3.609 (.605)***	1.804 (.666)***	1.932 (.831)**
Net wage partner/spouse	-.00009 (.00003)***	-.0003 (.00004)***	-.0003 (.00003)***	-.0005 (.00005)***
Age partner/spouse	.110 (.015)***	.117 (.021)***	.108 (.021)***	.114 (.028)***
Age sq. partner/spouse	-.001 (.0002)***	-.001 (.0003)***	-.001 (.0003)***	-.001 (.0004)***
Exp. partner/spouse	-.002 (.006)	-.010 (.010)	-.023 (.014)*	-.050 (.018)***
Exp. sq. partner/spouse	-.00009 (.0002)	.00005 (.0003)	.0008 (.0004)**	.002 (.0005)***
Educ. partner/spouse	.225 (.047)***	.206 (.074)***	-.007 (.088)	-.008 (.107)
Educ. sq. partner/spouse	-.009 (.002)***	-.008 (.003)**	-.0008 (.004)	-.001 (.004)
Region, where person lives (Western Germany)				
East-Germany	-.516 (.031)***	-1.032 (.060)***	-.515 (.032)***	-1.003 (.060)***
constant	-5.829 (.680)***	-4.687 (.990)***	.	.
time dummies, χ^2_{10} =	65.115***	115.718***	52.737***	96.926***
unobs. effects, χ^2_{17} =	.	.	573.04***	730.26***
LL	-17572.54	-13532.97	-17303.95	-13332.33
scale parameter ρ_a	.	.778 (.007)	.	.769 (0.007)

Source: GSOEP 1995-2005, own calculations. Different Probit specifications. 48,536 observations from 9,540 individuals. Standard errors in parenthesis: * significance at ten, ** at five, and *** at one percent. Year dummies are included in each procedure but not reported. a) Standard errors are robust to serial correlation in the individual scores across t; b) 24 points of quadrature; c) unobserved effects are specified as a linear projection on the (within) means of the regressors.

Table C.2: PARTICIPATION EQUATION, WOMEN, 1995-2005

	Pooled[a)]	Random Effects[b)]	Mundlak, pooled[a)c)]	Mundlak, R.E.[b)c)]
	(1)	(2)	(3)	(4)
Age	.078 (.040)**	.030 (.061)	.	.
Age square	-.0005 (.001)	.002 (.002)	.001 (.0002)***	.003 (.0004)***
Age triple	-1.00e-05 (7.81e-06)	-.00005 (1.00e-05)***	-.00003 (2.53e-06)***	-.00005 (4.23e-06)***
Education	.102 (.007)***	.239 (.012)***	.	.
Dummy Education	.018 (.032)	.023 (.041)	.	.
Foreigner	-.255 (.043)***	-.517 (.074)***	.	.
Health Sat.	.313 (.024)***	.276 (.032)***	.050 (.017)***	.120 (.035)***
Non labour inc.	-.0003 (.00002)***	-.0006 (.00002)***	-.0002 (.00002)***	-.0006 (.00002)***
Number of children				
up to 2 years old	-1.476 (.043)***	-2.557 (.063)***	-1.168 (.044)***	-2.268 (.065)***
between 3 - 5	-.827 (.029)***	-1.392 (.043)***	-.563 (.029)***	-1.154 (.046)***
between 6 - 16	-.361 (.017)***	-.566 (.025)***	-.187 (.018)***	-.385 (.030)***
Partner/Spouse variables				
Single	1.008 (.439)**	1.827 (.688)***	-.026 (.584)	.435 (.953)
Net wage partner/spouse	-.0002 (.00002)***	-.0002 (.00002)***	-.0001 (.00002)***	-.0002 (.00002)***
Age partner/spouse	.023 (.016)	.052 (.025)**	.002 (.022)	.036 (.037)
Age sq. partner/spouse	-.0003 (.0002)*	-.0005 (.0003)*	.0003 (.0002)	.0002 (.0004)
Exp. partner/spouse	.008 (.008)	-.002 (.013)	-.038 (.013)***	-.063 (.021)***
Exp. sq. partner/spouse	-.0002 (.0002)	-.0001 (.0003)	.0004 (.0002)	.0006 (.0004)
Educ. partner/spouse	.073 (.049)	.114 (.073)	-.002 (.059)	-.027 (.101)
Educ. sq. partner/spouse	-.002 (.002)	-.005 (.003)*	-.0005 (.002)	-.0009 (.004)
Region, where person lives (Western Germany)				
East-Germany	-.062 (.032)*	-.200 (.060)***	-.076 (.033)**	-.285 (.064)***
constant	-3.249 (.665)***	-4.355 (1.034)***	.	.
time dummies, χ^2_{10} =	51.25***	62.198***	46.025***	55.414***
unobs. effects, χ^2_{17} =	.	.	624.86***	760.78***
LL	-25488.93	-17217.55	-25226.81	-17027.29
scale parameter ρ_a	.	0.818 (.005)	.	.824 (.005)

Source: GSOEP 1995-2005, own calculations. Different Probit specifications. 48,763 observations from 10,081 persons. Standard errors in parenthesis: * significance at ten, ** at five, and *** at one percent. Year dummies are included in each procedure but not reported. a) Standard errors are robust to serial correlation in the individual scores across t; b) 24 points of quadrature; c) unobserved effects are specified as a linear projection on the (within) means of the regressors.

C.4 Summary statistics

Table C.3: SUMMARY, PARTICIPATION EQUATION, MEN, 1995-2005

	Entire Sample	Probit = 0	Probit = 1
Probit	.832 (.374)	0 (0)	1 (0)
Age	41.286 (10.997)	43.127 (13.310)	40.915 (10.431)
Age sq.	1825.462 (929.872)	2037.099 (1125.620)	1782.835 (879.096)
Age tr.	85498.640 (62658.360)	102698.600 (76567.650)	82034.300 (58860.570)
Education	12.206 (2.613)	11.253 (2.238)	12.398 (2.641)
Dummy educ.	.141 (.348)	.150 (.358)	.139 (.346)
Foreigner	.133 (.339)	.194 (.395)	.120 (.325)
Health sat.	2.566 (.414)	2.406 (.581)	2.598 (.363)
Non labour inc.	774.283 (970.113)	1438.180 (1054.049)	640.564 (894.576)
Number of children			
up to 2 years old	.082 (.288)	.056 (.240)	.087 (.297)
between 3 - 5	.118 (.353)	.074 (.289)	.127 (.364)
between 6 - 16	.480 (.816)	.351 (.750)	.506 (.827)
Partner/Spouse variables[a)]			
Single	.226 (.418)	.318 (.466)	.208 (.406)
Net wage partner/spouse	586.735 (637.155)	501.900 (649.150)	601.453 (633.907)
Age partner/spouse	40.648 (10.185)	44.220 (11.762)	40.028 (9.754)
Age sq. partner/spouse	1756.003 (856.123)	2093.697 (1017.939)	1697.415 (810.646)
Exp. partner/spouse	10.507 (9.176)	12.887 (11.168)	10.094 (8.719)
Exp. sq. partner/spouse	194.591 (299.486)	290.777 (394.703)	177.903 (276.306)
Educ. partner/spouse	11.732 (2.452)	11.024 (2.365)	11.855 (2.446)
Educ. sq. partner/spouse	143.648 (63.607)	127.112 (57.693)	146.517 (64.147)
Region, where person lives			
East-/West-Germany	.261 (.439)	.374 (.484)	.238 (.426)
N	48,536	8,137	40,399

Source: GSOEP 1995-2005, own calculations. All summary statistics are on individual-year level. Standard errors are in parenthesis.
a) The reported sample statistics for these variables are conditional on having a partner/ being married (Single = 0);

Table C.4: Summary, Participation Equation, Women, 1995-2005

	Entire Sample	Probit = 0	Probit = 1
Probit	.629 (.483)	0 (0)	1 (0)
Age	41.474 (11.194)	43.207 (12.232)	40.453 (10.400)
Age sq.	1845.396 (945.628)	2016.503 (1063.793)	1744.624 (852.646)
Age tr.	87019.930 (63840.140)	100022.100 (73899.680)	79362.450 (55690.880)
Education	11.911 (2.472)	11.182 (2.265)	12.340 (2.489)
Dummy Educ.	.127 (.333)	.126 (.332)	.127 (.333)
Foreigner	.130 (.336)	.192 (.394)	.093 (.291)
Health sat.	2.549 (.428)	2.490 (.497)	2.583 (.378)
Non labour inc.	802.090 (981.320)	1078.961 (1045.618)	639.030 (902.513)
Number of children			
up to 2 years old	.055 (.236)	.119 (.340)	.017 (.129)
between 3 - 5	.109 (.340)	.189 (.438)	.062 (.255)
between 6 - 16	.505 (.822)	.617 (.933)	.440 (.742)
Partner/Spouse variables[a)]			
Single	.214 (.410)	.153 (.360)	.250 (.433)
Net wage partner/spouse	1451.518 (1119.752)	1422.424 (1219.612)	1470.860 (1047.693)
Age partner/spouse	45.412 (11.295)	46.961 (12.267)	44.382 (10.474)
Age sq. partner/spouse	2189.811 (1054.607)	2355.833 (1173.289)	2079.437 (951.816)
Exp. partner/spouse	22.225 (11.327)	23.679 (11.961)	21.258 (10.777)
Exp. sq. partner/spouse	622.232 (528.356)	703.759 (583.731)	568.032 (480.496)
Educ. partner/spouse	12.039 (2.663)	11.673 (2.610)	12.283 (2.670)
Educ. sq. partner/spouse	152.038 (71.345)	143.073 (68.426)	157.997 (72.613)
Region, where person lives			
East-/West-Germany	.255 (.436)	.209 (.407)	.282 (.450)
N	48,763	18,074	30,689

Source: GSOEP 1995-2005, own calculations. All summary statistics are on individual-year level. Standard errors are in parenthesis.
a) The reported sample statistics for these variables are conditional on having a partner/being married (Single = 0);

Table C.5: Summary, Wage Equation, Men, 1995-2005

	Mean	Std. dev.	10% pctl.	90% pctl.
Log hourly wage	2.578	.407	2.071	3.093
Health sat.	2.599	.360	2.095	2.893
Age	41.020	10.298	28	56
Age sq.	1788.715	870.012	784	3136
Age tr.	82257.240	58301.720	21952	175616
Unempl. exp.	.380	1.056	0	1.100
Unempl. exp. sq.	1.258	8.746	0	1.210
Firm tenure	11.314	10.052	1.100	27
Firm tenure sq.	229.057	348.743	1.210	729
Education	12.418	2.642	10.500	18
Dummy educ.	.142	.349	0	1
Part-time	.018	.135	0	0
Foreigner	.119	.324	0	1
Lg. unempl. (fed. st.)	12.768	.569	12.150	13.630
Lg. vac. (fed. st.)	10.443	.839	9.136	11.404
Firm size (<20 employees)$^{a)}$				
20 - 199	.301	.459	0	1
200 - 1999	.237	.425	0	1
$\geq$ 2000 employees	.262	.440	0	1
Firm size miss.	.017	.127	0	0
Region, where person works (Western Germany)				
Eastern Germany	.223	.416	0	1
Occupation Dummies (High Service)				
Low Service	.185	.388	0	1
Routine Non-Manual	.041	.198	0	0
Skilled Manual	.308	.462	0	1
Semi-unskilled Manual	.211	.408	0	1
Farm Labour	.011	.106	0	0
Missing occ.	.086	.280	0	0
Sector Dummies (Agr., forestry, fishing)				
Unknown sector	.022	.147	0	0
Energy, water, mining	.015	.123	0	0
Manufacturing	.369	.483	0	1
Construction	.111	.315	0	1
Trade	.086	.280	0	0
Transport, communication	.042	.200	0	0
Financial serv., insurance	.024	.154	0	0
Other services	.089	.285	0	0
State	.229	.420	0	1
Instruments				
Num. vis. doc. (last 3 months)	1.759	3.316	0	4
Days off due to illness $(t-1)^{b)}$	8.951	21.365	0	21
Non labour inc.	629.209	879.454	0	1711.065
Single	.203	.402	0	1
Net wage partner/spouse$^{c)}$	603.609	633.797	0	1450.677
Age partner/spouse	40.027	9.667	28	53
Age sq. partner/spouse	1695.632	802.981	784	2809
Exp. partner/spouse	10.109	8.686	.700	23.500
Exp. sq. partner/spouse	177.640	274.742	.490	552.250
Educ. partner/spouse	11.867	2.446	9	15
Educ. sq. partner/spouse	146.809	64.209	81	225

Source: GSOEP 1995-2005, own calculations. All summary statistics are on individual-year level (39,048 observations). Persons with participation in only one year and individuals with missing wages are dropped from the sample. *a)* For dummy variables, the basis categories are given in parenthesis; *b)* the reported sample statistics is conditional on whether the person was working last year. The variable is set to zero otherwise; *c)* the reported sample statistics for these variables are conditional on having a partner/ being married (Single = 0).

Table C.6: Summary, Wage Equation, Women, 1995-2005

	Mean	Std. Dev.	10% pctl.	90% pctl.
Log horuly wage	2.362	.400	1.839	2.834
Health sat.	2.584	.376	2.095	2.893
Age	40.610	10.264	26	55
Age sq.	1754.498	843.538	676	3025
Age tr.	79830.600	55148	17576	166375
Unempl. exp.	.449	1.105	0	1.400
Unempl. exp. sq.	1.423	10.536	0	1.960
Firm tenure	9.405	8.647	1	23.200
Firm tenure sq.	163.230	270.144	1	538.240
Education	12.360	2.488	10	16
Dummy Educ.	.129	.336	0	1
Part-time	.367	.482	0	1
Foreigner	.092	.289	0	0
Lg. unempl. (fed. st.)	12.752	.566	12.150	13.626
Lg. vac. (fed. st.)	10.367	.861	9.060	11.404
Firm size (<20 employees)$^{a)}$				
20 - 199	.295	.456	0	1
200 - 1999	.228	.420	0	1
≥ 2000 employees	.200	.400	0	1
Firm size miss.	.018	.132	0	0
Region, where person works (Western Germany)				
Eastern Germany	.275	.447	0	1
Occupation Dummies (High Service)				
Low Service	.259	.438	0	1
Routine Non-Manual	.202	.402	0	1
Skilled Manual	.068	.252	0	0
Semi-unskilled Manual	.172	.377	0	1
Farm Labour	.009	.093	0	0
Missing occ.	.219	.414	0	1
Sector Dummies (Agr., forestry, fishing)				
Unknown sector	.023	.149	0	0
Energy, water, mining	.004	.061	0	0
Manufacturing	.171	.376	0	1
Construction	.017	.130	0	0
Trade	.154	.361	0	1
Transport, communication	.023	.149	0	0
Financial serv., insurance	.031	.174	0	0
Other services	.200	.400	0	1
State	.370	.483	0	1
Instruments				
Num. vis. doc. (last 3 months)	2.382	3.470	0	5
Days off due to illness $(t-1)^{b)}$	9.567	22.253	0	21
Non labour inc.	629.226	890.755	0	1693.780
Single	.247	.431	0	1
Net wage partner/spouse$^{c)}$	1472.655	1045.818	0	2636.535
Age partner/spouse	44.476	10.391	31	59
Age sq. partner/spouse	2086.081	945.872	961	3481
Exp. partner/spouse	21.349	10.705	6.900	36
Exp. sq. partner/spouse	570.349	477.907	47.610	1296
Educ. partner/spouse	12.297	2.673	10	18
Educ. sq. partner/spouse	158.374	72.758	100	324

Source: GSOEP 1995-2005, own calculations. All summary statistics are on individual-year level (29,304 observations). Persons with participation in only one year and individuals with missing wages are dropped from the sample. *a)* For dummy variables, the basis categories are given in parenthesis; *b)* the reported sample statistics is conditional on whether the person was working last year. The variable is set to zero otherwise; *c)* the reported sample statistics for these variables are conditional on having a partner/being married (Single = 0).

References

Chamberlain, Gary, "Analysis of Covariance with Qualitative Data," *Review of Economic Studies*, 1980, *47*, 225–238.

____ , "Panel data," *in Zvi Griliches and Michael D. Intriligator (eds), Handbook of Econometrics*, 1984, *Volume 2*, 1247–1318.

Contoyannis, Paul and Nigel Rice, "The Impact of Health on Wages: Evidence from the British Household Panel Survey," *Empirical Economics*, 2001, *26*, 599–622.

____ **, Andrew Michael Jones, and Nigel Rice**, "The dynamics of health in the British Household Panel Survey," *Journal of Applied Econometrics*, 2004, *19* (4), 473–503.

Dustmann, Christian and Maria Engracia Rochina-Barrachina, "Selection correction in panel data models: an application to labour supply and wages," *IZA Discussion Paper 162*, 2000.

Gambin, Lynn M., "The Impact of Health on Wages in Europe – Does gender matter?," *HEDG Working Paper*, June 2005, *03.*

Gordo, Laura Romeu, "Effects of short- and long-term unemployment on health satisfaction *evidence from German data," *Applied Economics*, 2006, *38* (20), 23352350.

Grossman, Michael, "The Human Capital Model," in Anthony J. Culyer and Joseph P. Newhouse, eds., *Handbook of Health Economics*, Vol. 1A, Amsterdam: Elsevier Science B.V., 2001, pp. 347–409.

Halliday, Timothy J. and John A. Burns, "Heterogeneity, State Dependence and Health," *University of Hawaii at Manao, Department of Economics Working Paper Series*, 2005, *3.*

Haveman, Robert, Barbara Wolfe, Brent Kreider, and Mark Stone, "Market Work, Wages, and Men's Health," *Journal of Health Economics*, 1994, *13*, 163–182.

Heckman, James J., "Dummy Endogenous Variables in a Simultaneous Equation System," *Econometrica*, 1978, *46* (4), 931–60.

Kyriazidou, Ekaterini, "Estimation of a panel data sample selection model," *Econometrica*, 1997, *55*, 1335–1364.

Lee, Lung-Fei, "Health and Wage: A Simultaneous Equation Model with Multiple Discrete Indicators," *International Economic Review*, February 1982, *23* (1), 199–221.

Maddala, G.S., *Limited-Dependent and Qualitative Variables in Econometrics*, Cambridge: Cambridge University Press, 1983.

Mincer, Jacob, "Investment in Human Capital and Personal Income Distribution," *Journal of Political Economy*, August 1958, *66* (4), 281–302.

———, "Schooling, Experience and Earnings," *New York: National Bureau of Economic Research*, 1974.

Mundlak, Yair, "On the pooling of time series and cross section data," *Economatrica*, 1978, *46*, 69–85.

Rochina-Barrachina, Maria Engracia, "A new Estimator for Panel Data Sample Selection Models," *Annales d'Economie et de Statistique*, 1999, *55/56*, 153.181.

Semykina, Anastasia and Jeffrey M. Wooldridge, "Estimating Panel Data Models in the Presence of Endogeneity and Selection: Theory and Application," *mimeo*, October 2005.

SOEP Group, "The German Socio-Economic Panel (GSOEP) after more than 15 years - Overview," in Elke, Holst, Dean Lillard, and Thomas A. DiPrete, eds., *Proceedings of the 2000 Fourth International Conference of German Socio-Economic Panel Study Users (GSOEP2000)*, Vierteljahreshefte zur Wirtschaftsforschung, 2001, 70(1), pp. 7-14.

Verbeek, Marno and Theo Nijman, "Testing for Selectivity Bias in Panel Data Models," *International Economic Review*, 1992, *33*, 681–703.

Wooldridge, Jeffrey M., "Selection correction for panel data models under conditional mean independence assumption," *Journal of Econometrics*, 1995, *68*, 115–132.

———, *Econometric analysis of cross section and panel data*, Cambridge and London: MIT Press, 2002.

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